



Essays on the Economics of Wage Inequality

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Essays on the Economics of Wage Inequality

A dissertation presented by

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to

The Committee for the PhD in Business Economics

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Essays on the Economics of Wage Inequality

ABSTRACT

In this dissertation I examine changes in wage inequality in two chapters. In the first chapter, I examine the slowdown in the relative demand for college-educated labor in the U.S. since the early 1980s. A large literature suggests that this puzzling slowdown is primarily the result of non-monotone changes in the demand for skill, particularly since the mid-1990s, induced by the introduction of computers to the labor market. In these two essays, I develop a complementary result: I show that roughly 10-60% of the gap in the annual growth rates of the relative demand for college-educated workers between the 1963-1982 and 1982-2008 periods can be closed by adjusting for shifts in supply and demand within schooling groups; however, a slowdown in relative demand growth beginning in 1993, well-documented in the literature and potentially-related to recent technological changes, remains pronounced across all specifications.

In the second chapter, I examine changes in relative wages in Brazil. From 1995 to 2002, per capita GDP and schooling wage differentials in Brazil remained relatively stable; during the boom in international commodity prices from 2002 to 2011, per capita GDP grew a robust 2.7% annually while wage differentials along all points of the schooling distribution compressed dramatically. In contrast to a recent literature that interprets these patterns as evidence of trend changes in relative labor demand favoring less-educated workers, I show that a standard two schooling group model of the labor market appears inconsistent with observed movements of

relative wages and supplies, potentially implying that demand shifts among male workers in Brazil from 1995 to 2011 were heterogeneous within broad schooling groups. As an alternative to grouping workers solely by their level of schooling, I propose an alternative model that compares three types of workers: low-skill (less-educated, less-experienced), high-skill (more-educated, more-experienced), and middle-skill (either more-educated or more-experienced, but not both).

Contents

Abstract	iii
Acknowledgments	vi
I. How Much has Growth in the Relative Demand for College-Educated Workers Slowed Since the Early 1980s?	1
I.A Introduction	1
I.B A Simple Generalization of the Standard Model.	9
I.C Data and Empirical Measures	18
I.D Changes in the Relative Demand for Potential Experience.	23
I.E Changes in the Relative Demand for Female Labor.	53
I.F Changes in the Relative Demand for the College-Educated	73
I.G Conclusion	95
II. Did Relative Labor Demand Shift Towards Less-Educated Workers in Brazil, 1995-2011?	96
II.A Introduction	96
II.B Data and Empirical Measures	98
II.C Patterns of Wage Inequality, Relative Wages and Supplies.	101
II.D Modeling Relative Wages in Brazil, 1995-2011	119
II.E Conclusion	136
III. References	138
IV. Appendices	144

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I How Much has Growth in the Relative Demand for College-Educated Workers Slowed Since the Early 1980s?

I.A Introduction

A large literature argues that a long-run trend in relative demand favoring college- relative to high school-educated labor has played an important role in changes in U.S. wage inequality over the past century, and further that this relative demand trend reflects skill-biased technological change, as modeled and measured in the "standard" supply and demand model with two schooling groups (Goldin and Katz, 2008; 2009; Katz and Murphy, 1992). A closely-related literature notes that the widespread use of information technology beginning in the early 1980s may be an observable and important driver of skill-biased technological change (Bound and Johnson, 1992; Krueger, 1993). However, rather than accelerating, growth in the relative demand for college-educated labor, as measured in the standard model, appears to have remained steady or to have decreased slightly as personal computer use became widespread in the early 1980s,¹ and appears to have sharply slowed beginning in the mid-1990s as observable progress in information technology continued apace (Autor et al., 1998; Goldin and Katz, 2009). A prominent explanation for these puzzling patterns is that while information technology is indeed a relative complement for the most-skilled workers in the labor market, this technology may have, especially since the mid-1990s, depressed demand for the routine tasks performed by "middle-skilled" workers - including many workers with college experience - leaving demand for the manual tasks performed mainly by workers with no more than a high school degree relatively untouched (Autor et al., 2003; 2006; 2008; 2013; Levy and Murnane, 2004; 2006; Weiss, 2008).²

In this paper, rather than seeking to understand the changes in technology that potentially undergird the slowdown of relative demand for college-educated labor since the early 1980s, I re-examine the measurement of the slowdown itself. I return to the standard two schooling group model to examine

¹During the 1976-1981 period, the standard model performs (relatively) poorly in predicting the time series of the college wage premium, potentially as a result of non-market factors (such as the union-negotiated wage settlements of the 1970s) that complicate the measurement of relative demand and growth across periods that include this range of earnings years (Goldin and Katz, 2009). However, annual growth in the relative demand for the college-educated during the 1982-1993 period is consistently similar to, or slower than, growth during the 1963-1975 and 1963-1982 periods across a wide range of specifications. I discuss the measurement of relative demand growth in detail in the paper.

²A number of authors have also interpreted the measured slowdown in the relative demand for college-educated labor as evidence against the usefulness of the supply and demand approach to understanding changes in the wage structure (Beaudry and Green, 2005; Card and DiNardo, 2002; Lemieux, 2006).

whether, from 1963 to 2008, shifts in labor demand *within* college- and high school-educated workers, including large trend breaks in relative labor demand favoring women concurrent with the beginning of the computer revolution, confound the measurement of relative demand *between* these schooling groups. To illustrate why within-schooling group demand shifts may be problematic, consider a simple example. Averaged across the 1963-2008 earnings years, women represented 35% and 41% of workers with at least a college degree, and exactly a high school degree, respectively.³ Because, on average within either of these schooling groups, women represent far fewer than half of all workers, a shift in labor demand favoring women within a schooling group will, all else equal, tend to decrease the average (fixed employment-weighted) wage of that schooling group. Since women are, on average, less-represented among the college-educated than among the high school-educated, identical shifts in relative labor demand favoring women within both of these schooling groups will tend to decrease the average wage of the college-educated more than that of the high school-educated; moreover, in my empirical work, I find evidence suggesting that trend breaks in the relative demand for women since the late 1970s are larger among the college-educated than among the high school-educated, further decreasing the traditional fixed-weight measure of the college wage premium. Since, in the standard model, relative demand for the college-educated is measured by comparing increases in the college wage premium to movements in relative supply, failing to account for demand shifts within schooling groups that pull down the college wage premium will tend to understate true relative demand growth.⁴

While demand within college- and high school-educated workers may vary across a wide range of worker characteristics,⁵ in this paper I consider variation across two specific traits: labor market experience and gender. I begin by introducing a simple generalization of the standard two schooling group model that allows demand to vary within both college- and high school-educated labor across

³ Among a sample of all employed workers who possess between 0 and 39 years of (Mincerian) potential labor market experience, and measured as total usual hours worked. I discuss the measurement of employment shares in detail in the empirical sections of the paper.

⁴ The hypothesis explored in this paper is related to, but distinct from, the hypothesis that changes in cohort quality (rather than shifts in demand) have depressed the college wage premium over time. Evidence supporting this hypothesis is mixed: while estimates of Carneiro and Lee (2011) suggest that (traditionally-unmeasured) declines in the quality of college graduates resulted in a 6% decline in the college wage premium from 1960 to 2000, other authors conclude that cohort quality changes play a much smaller role in driving changes in educational differentials (Juhn et al., 1993; Juhn et al., 2005; Katz and Murphy, 1992).

⁵ For example, in the last two decades, substantial wage polarization within the two schooling groups of the standard model has occurred along the *schooling* margin: for example, during this period the returns to post-graduate education have increased faster than the returns to college (Autor et al. 2008; Goldin & Katz, 2007). In this paper, I abstract from these phenomena.

these demographic dimensions. The key wage equation of the model suggests the importance of three separate sets of supply and demand forces which I investigate, in sequence, using three separate sets of worker comparisons. First, I compare workers with the same level of schooling and of the same gender, but with different levels of (Mincerian) potential labor market experience. Second, I compare workers with the same level of schooling, but of different genders. Finally, I perform the primary comparison of interest: college- relative to high school-educated workers. For each comparison, observed patterns of wages and supplies, combined with an elasticity of substitution between compared groups, yield estimated shifts in relative demand. The goal of the paper is to understand how shifts in supply and demand within schooling groups affect the measurement of the relative demand for college-educated labor measured in the final comparison. Within each step, I slowly impose the structure of the model: I begin with simple reduced-form comparisons of the comovement of relative wages and supplies for clues to the potential role of relative demand shifts, then discipline these comparisons by applying the model across a broad range of parameter values, and finally - when considering changes in the relative demand for potential experience and the relative demand for college-educated labor - I estimate particular values of model parameters. In each comparison, I incorporate the supply and demand effects measured in the preceding steps.

I.A.1 Changes in the Relative Demand for Potential Experience

In the first section, I estimate the shifts in supply and demand potentially-responsible for changes in the returns to experience since the early 1960s. My analysis addresses two divisions in the literature. First, while a set of papers suggests a large role for steadily-increasing relative demand for more-experienced workers among males since the early 1960s (Juhn et al., 1993; Katz and Murphy, 1992; Murphy and Welch, 1992), recent generalizations of the standard model have tended to examine the hypothesis that supply shifts alone can explain changes in the return to experience among men (Autor et al., 2008; Card and Lemieux, 2001). Second, despite substantial evidence of large increases in the returns to (actual) labor market experience among women beginning in the late 1970s⁶ (Blau and Kahn, 1994; 1997; O'Neill and Polachek, 1993) - a different pattern than observed among men - the bulk of the literature examining changes in U.S. experience differentials, including each paper referenced in the

⁶Evidence on the *causes* of this increase in the return to labor market experience among women is limited; O'Neill and Polachek suggest that declines in labor market discrimination, widely hypothesized to play a role in the closing of the gender wage gap since the late 1970s (Becker, 1985), may have played a demand-side role if employers became more-willing to train and promote women.

previous sentence, either pools male and female workers of the same potential experience category, or focuses exclusively on male workers when making these comparisons.⁷ In my model I allow demand, and not only supply, to influence experience differentials, and because I separately-examine workers with the same level of schooling and of the same gender, I am able to uncover differences between men and women in the supply and demand forces responsible for changes over time in the returns to potential labor market experience.

I first examine the simple comovement of wages and supplies over longer-run (21-year) periods across potential experience groups. Among men, relative wages and supplies have moved in opposite directions (e.g., the returns to experience within both college- and high school-educated men rose as the baby boom cohort of younger, less-experienced men entered the labor market from 1964-1985), while among women they have tended to move together (e.g., the returns to experience among both college- and high school-educated women rose from 1985-2006 despite concurrent increases in the relative supply of more-experienced women), potentially-implying an important role for longer-run demand shifts among women. In order to investigate the key parameter of interest - the elasticity of substitution between workers with different levels of potential experience among workers of the same schooling and gender - I explore in detail the pioneering empirical approach of Card and Lemieux (2001), who examine variation in the time series of the college wage premium across potential experience groups. Here I highlight two findings: first, for a wide range of values of this key parameter within the model, the hypothesis that all differences in the time series of the college wage premium across potential experience groups are explained solely by supply shifts cannot be rejected for male workers, but can be rejected among female workers, again suggesting a potential role for relative demand shifts. Second, I find evidence that, among both male and female workers, the elasticity of substitution between potential experience groups is approximately 5.025, which is very close to the estimates of Card and Lemieux, all of which are close to 5.

With this estimate in hand, I then fully-apply the model to estimate supply and demand shifts across potential experience groups. Consistent with the literature from the early 1990s referenced above, the fitted model implies a smooth and consistent trend in demand across the entire 1964-2006 period favoring the more-experienced among college-educated males. By contrast, within high school-

⁷This may be due, in part, to difficulties in measuring actual labor market experience among women using a Mincerian potential experience variable (the best available experience measure in the Current Population Survey, but a function of only age and schooling) across a period of significant increase in female labor market participation (O'Neill, 2003a; 2003b). I return to this point below.

educated male workers and both high school- and college-educated women, I find evidence of a trend break in relative demand, beginning in the late 1970s, favoring more-experienced workers. Consistent with work examining the closing of the gender wage gap during the 1980s (Blau and Kahn, 1994; 1997; O'Neill and Polachek, 1993), shifts in relative demand among women are particularly rapid;⁸ moreover, demand shifts appear to play a larger role in determining changes in potential experience differentials for women: on average, slightly more than half of the (gross) movement in these wage differentials is accounted for by relative demand shifts among women, compared to slightly less than half among men. Finally, I show that mis-measurement of actual labor market experience among women appears unlikely to explain these results.

I.A.2 Changes in the Relative Demand for Female Labor

I incorporate these results into the second empirical section of the paper, in which I examine shifts in supply and demand potentially-responsible for the evolution of gender wage gaps among both college- and high school-educated workers over the 1963-2008 earnings years. In contrast to the literature on the returns to labor market experience, empirical work on changes over time in gender wage differentials has more-consistently suggested a large role for labor demand. Shifts over time in labor demand favoring female workers appear to help explain why, in spite of large increases in female labor force participation, gender differentials were relatively stable among full-time workers from 1930 to 1980 (Goldin, 1990)⁹ and have narrowed since the late 1970s (Blau and Kahn, 1994; 1997; 2004; Katz and Murphy, 1992; O'Neill and Polachek, 1993; Smith and Ward, 1989), potentially as a result of demand shifts at the industry-by-occupation level such as the decline of male-intensive basic manufacturing work and production jobs (Katz and Murphy, 1992, see Table 5; Blau and Kahn, 1997), or more difficult to measure, but potentially-important forces such as a decline in labor market discrimination (Becker, 1985) or an increase in the relative demand for "brains" relative to "brawn" during the 1980s (Beaudry and Lewis, 2014; Galor and Weil, 1996; Weinberg, 2000; Welch, 2000).

⁸For example, among college-educated women, the demand for prime-age workers (workers with between 20-29 years of potential experience) increased roughly 3.5% faster per year than the leave-out average among all college-educated women, as measured by a standard relative demand index.

⁹For example, the simple supply and demand framework of Goldin (1990; Chapter 5) suggests that increases in the relative demand for female labor account for more than half of the increase in female labor force participation among urban white married women from 1940 to 1980, during which period the relative female wage was steady at roughly 60% of male wages. This supply and demand analysis is to be distinguished, however, from earlier analyses which consider only shifts in labor supply (Mincer, 1962; Smith and Ward, 1984; 1985).

I begin by investigating simple patterns of comovement between the relative wages and supplies of female labor within both college- and high school-educated labor. This reduced-form evidence provides two preliminary suggestions about the role of labor demand in determining gender wage differentials. First, consistent with the literature cited above, within both college- and high school-educated workers, a compression of the log gender wage gap by roughly 20% between 1978 and 1993 despite large simultaneous increases in the relative supply of female labor - a joint rise in both relative quantity and price - suggests a potentially-important role for increases in the relative demand for female labor beginning in the late 1970s. Second, while the relative supply of female labor among the college-educated saw strong, nearly-linear growth across the entire 1963-2008 period, rapid increase in the relative supply of female labor within the high school-educated abruptly halted in 1982, concurrent with the well-documented slowdown in the educational attainment of the labor force in the same year (Katz and Murphy, 1992).¹⁰ This second fact hints that shifts in relative demand favoring female workers may have been larger among the college-educated than among the high school-educated since, all else equal, the continuous increase in the relative supply of women among the college-educated may have acted to increase the gender wage gap among college-educated workers, while this upward pressure on the gender wage gap among the high school-educated was relieved in 1982.¹¹

To more-formally investigate these patterns, I use the model, together with observed movements in relative wages and supplies, to investigate shifts in relative demand. Because variation in the relative supply of female labor is likely to be driven in large part by changes along the labor force *participation* margin and is thus, even in the short run, not likely to be orthogonal to changes in the relative demand for female labor, I do not attempt (as in the previous, or the following, section) to estimate the relevant elasticity of substitution. Instead, I investigate shifts in relative demand for female labor implied by a wide range of choices for the elasticity of substitution between female and male labor suggested by existing estimates in the literature, which range from 1.5 to 4.5 and are most-commonly

¹⁰ As has been well-documented (e.g., Goldin and Katz, 2008, Figure 7.1), the slowdown in the college graduation rate in the U.S. is much more-pronounced among males.

¹¹ An alternative hypothesis is that men and women may not be direct substitutes within both high school and college labor. For example, a number of authors model all female labor as a direct substitute for only low-skilled male labor (Grant and Hamermesh, 1981; Topel, 1994; 1997). However, in this paper I argue that, since the skills gained by attending college are of primary importance in the post-World War II labor market (Becker, 1964; Goldin and Katz, 2008; Mincer, 1974), men and women within both the college- and high school-educated may be properly-modeled as direct substitutes in production, a notion supported by recent empirical evidence (Acemoglu and Autor, 2004; Blau and Kahn, 1997; Juhn and Kim, 1999) discussed in the paper.

found to be around 2.5 (Acemoglu and Autor, 2004; Blau and Kahn, 2004; Layard, 1982; Lewis, 1985; Weinberg, 2000). Across this wide range of parameter values, the fitted model consistently suggests two conclusions. First, the relative demand for female labor rose dramatically over the 1978-1993 period: for example, a value for the elasticity of substitution between female and male labor of 2.5 implies that the relative demand for female labor within college- and high school-educated workers across this period increased by roughly 86% and 67%, respectively, as measured by a standard relative demand index. Second, for every elasticity value under consideration, demand shifts favoring women were larger among the high school-educated than among the college-educated from 1963-1978, but this trend reversed in 1978: over the next 30 years, increases in the relative demand for women were larger among the college-educated than among the high school-educated. In the last part of this section, I discuss (i) differences between these two main empirical findings and those of the existing literature, and (ii) the alternative hypothesis that these patterns are driven by unobserved changes in cohort quality across time.

I.A.3 Changes in the Relative Demand for the College-Educated

These two key empirical findings, together with the supply and demand shifts between potential experience groups estimated in the previous section, have important implications for the measurement of changes in the relative demand for college-educated labor, which I investigate in the third and final empirical section of the paper. As noted above, a key puzzle in this area is that, according to the standard model, the relative demand for college-educated labor has slowed since information technology became widespread in the labor market in the early 1980s¹² (Acemoglu and Autor, 2011; Autor et al., 1998; 2008; Card and DiNardo, 2002; Goldin and Katz, 2008; 2009). In order to investigate the key hypothesis of the paper - that shifts in supply and demand within college- and high school-educated workers affect the measurement of this slowdown - I begin by deriving a generalization of the standard Katz and Murphy (1992) wage equation that relates the relative wage and supply of the college-educated. I use this equation to make two simple theoretical points. First, in the general case of the model, standard fixed-weight measures of the relative wage and supply of college-educated labor must first be adjusted for shifts in supply and demand *within* schooling groups before proceeding with an analysis of shifts in supply and demand *between* schooling groups. I then show that, in the

¹²Card and DiNardo (2002) note that while the personal computer had been previously commercially introduced, widespread use of computer technology is generally associated with the introduction of the IBM-PC in 1981.

general model, the relative demand for college-educated labor equals the standard measure plus a set of adjustments for these supply and demand shifts that can be explicitly derived from the estimates of previous sections of the paper.

I begin the empirical analysis with simple comparisons of the movement of (adjusted) relative wages and supplies. First, I find that, across the 1963-2008 earnings years, the time series of the adjusted relative supply measures implied by the general model are nearly-identical to the standard fixed-weight measure. By contrast, around the introduction of the personal computer, growth in the adjusted college wage premium begins to significantly-outpace growth in the standard fixed-weight measure: across the 1982-2008 earnings years, within-schooling group shifts in supply and demand pushed down the log college wage premium by roughly 4-12%. Taken together, these patterns suggest that, after adjusting for shifts in supply and demand within schooling groups, growth in the relative demand for college-educated after 1982 may be more-rapid than implied by standard estimates.

I confirm this suggestion by explicitly calculating, for a wide range of parameter values within the general model, growth in the relative demand for college-educated labor. My analysis highlights four key differences between the standard measure of relative demand growth and general measures, which adjust for within-schooling group shifts in supply and demand. First, across nearly all specifications and time periods, growth in the relative demand for college-educated workers is more-rapid after adjusting for shifts in supply and demand within schooling groups. Second, differences between general and standard measures are most-prominent after 1982, and my preferred estimates suggest that roughly 10-60% of the gap in the annual growth rates of the (standard measure of) relative demand for college-educated workers between the 1963-1982 and 1982-2008 periods can be closed by adjusting for shifts in supply and demand within schooling groups. Third, these differences are driven by two shifts in relative demand within schooling groups: shifts in relative demand favoring women, and favoring those workers with more potential labor market experience among women, that began in the late 1970s and were investigated in previous sections of the paper. Finally, although adjustments for shifts in supply and demand within schooling groups appear to increase the measured pace of growth in the relative demand for the college-educated after 1982, they do not decrease the well-documented slowdown in relative demand beginning in 1993, which persists across all specifications of the general model.

Taken together, evidence presented in this paper is complementary with work that suggests that the introduction of information technology to the labor market since the mid-1990s slowed down

the relative demand for college-educated workers by depressing the demand for specific tasks, some of which are performed by the college-educated (Autor et al., 2003; 2006; 2008; 2013; Levy and Murnane, 2004; 2006; Weiss, 2008). In particular, estimates in this paper suggest that the relative demand for college-educated workers may have decreased less than previously-thought in the years after the introduction of the IBM-PC in 1981, but appear to confirm the economic importance of the technological shifts responsible for the sharp slowdown in the relative demand for college-educated labor since the mid-1990s. The remainder of the paper is organized as follows. In Section II, I outline a simple generalization of the standard two schooling group model of the labor market. In Section III, I outline the data, basic processing procedures, and the empirical measures of wages, supplies and relative demands that I use in the paper. In Sections IV and V, I estimate shifts in relative demand between potential experience and gender groups, respectively. Then, in Section VI I examine observed changes in wages and supplies of college- and high school-educated workers and I show how the supply and demand effects estimated in Sections IV and V affect the measurement of the relative demand for college-educated labor. Section VII concludes.

1.B A Simple Generalization of the Standard Model

In this section, I outline a simple generalization of the "standard" supply-and-demand model with two schooling groups, pioneered by Katz and Murphy (1992).¹³ While a number of papers propose models that mathematically nest the standard two schooling group framework, the current model departs from the bulk of this literature in two ways. First, demand is allowed to vary within, rather than solely between, schooling groups - a generalization I refer to as relaxing the assumption of "common demand shifts." Second, I assume a structure of aggregate production that emphasizes comparisons of workers both within and between gender. These two departures allow me to examine how shifts in both (observed) relative supply and (estimated) relative demand within both college- and high school-educated workers affect the measurement of the relative demand for college-educated labor in the standard model. After describing the model I derive the key wage equation, which implies that all heterogeneity in the time series of wages is due to three sets of supply and demand forces investigable with three separate worker comparisons explored in Sections IV, V and VI of this paper: comparisons

¹³While Freeman (1976, Appendix B) estimates equations similar to (but containing important differences from) those derived from the two schooling group model, Katz and Murphy lay out the first general framework of the supply and demand for workers of different demographic groups. The two schooling group special case has proven especially relevant empirically, and is generally referred to as the "standard" or "canonical" model (e.g., Acemoglu and Autor, 2011).

across potential experience, gender, and schooling dimensions, respectively. Finally, using the key wage equation as a guide, I discuss how each set of worker comparisons suggested by the model relates to existing literatures on changes in experience, gender and schooling wage differentials.

I.B.1 The Assumed Structure of Production and the Key Wage Equation

I follow the approach of Katz and Murphy and the subsequent literature in modeling the wages and supplies of, and the demands for workers of two schooling types. First, I examine the wages of workers of two distinct types: those who complete exactly 12 years of schooling before labor market entry (henceforth "high school" labor), and those who complete 16 or more years of schooling before labor market entry (henceforth "college" labor). I then assume that the supply of these two worker types may be captured by two schooling "equivalents:" amalgams of workers of all levels of schooling, organized into either the college or high school category. I use this approach to wages, supplies and demands within all demographic groups described below.

I assume that the production of aggregate output Y can be approximated by a CES function with two factors - college equivalent and high school equivalent labor:

$$(1) \quad Y_t = \left[(A_{Ct}C_t)^{\frac{\sigma_s-1}{\sigma_s}} + (A_{Ht}H_t)^{\frac{\sigma_s-1}{\sigma_s}} \right]^{\frac{\sigma_s}{\sigma_s-1}}$$

Here C_t and H_t are indices that capture the effective quantity of labor supplied by college and high school equivalents in period t , A_{Ct} and A_{Ht} capture labor-augmenting technological change common to all workers within these schooling groups, and σ_s is the aggregate elasticity of substitution between these two factors, and is assumed to be (strongly) positive.¹⁴ I introduce the possibility of gender-specific shifts in supply and demand within both college and high-school equivalent labor by assuming that, for each schooling group $s \in \{C, H\}$:

¹⁴For simplicity, equation (52) as well as equations (2) and (3) below assume that all technological shifts are due solely to labor-augmenting technological change. By contrast, some authors assume that these shifts are due exclusively to changes in either share parameters (Goldin and Katz, 2008; 2009), or to both labor-augmenting and share parameters (Autor et al., 1998). Below, I show that the standard relative demand indices I employ in this paper are consistent with all three classes of assumption.

$$(2) \quad s_t = \left[(A_{sFt} F_{st})^{\frac{\sigma_g - 1}{\sigma_g}} + (A_{sMt} M_{st})^{\frac{\sigma_g - 1}{\sigma_g}} \right]^{\frac{\sigma_g}{\sigma_g - 1}}$$

where F_{st} and M_{st} are indices that represent the effective quantity of equivalent labor supplied by women and men, respectively, within schooling group s at time t . The parameter σ_g captures the elasticity of substitution between genders, and is assumed to be strongly greater than zero and identical for both schooling groups. Finally, I follow Card and Lemieux (2001) and many subsequent authors in defining sub-aggregates that distinguish between workers with different levels of (Mincerian) potential labor market experience. I assume that, for each schooling group $s \in \{C, H\}$, and for each gender $g \in \{F, M\}$:

$$(3) \quad g_{st} = \left[\sum_{e=1}^E (A_{sget} \pi_{sge} g_{set})^{\frac{\sigma_e - 1}{\sigma_e}} \right]^{\frac{\sigma_e}{\sigma_e - 1}}$$

where g_{set} represents the actual (not effective) quantity of equivalent labor supplied by workers of schooling type s , gender g and potential experience e , π_{sge} represents the quantity of relative efficiency units of labor supplied in every period by an equivalent worker in that group, and E represents the total number of potential experience groups.¹⁵ The parameter σ_e captures the elasticity of substitution between potential experience groups within each schooling-by-gender aggregate, and is assumed to be (strongly) positive. Finally, the assumption of a competitive labor market ensures that observed prices and quantities are set "on the demand curve," and together with the assumed structure of aggregate production in equations (52), (2) and (3), completely-describes the model.

Relative to common variants of the standard framework, two features distinguish the model. First, I introduce the demand shifters A_{sgt} , which capture differences in labor demand between males and females within each schooling group, and the demand shifters A_{sget} , which capture differences in labor demand between more- and less-experienced workers within a given schooling-by-gender group. The assumption of "common demand shifts" - that $A_{sgt} = 1$ for each s , g and t and $A_{sget} = 1$ for each

¹⁵ To simplify notation, I do not explicitly distinguish between the (common) effective quantity of labor supplied in each period by all equivalent workers and the (relative) effective quantities that are specific to each individual demographic group. Formally, I can write $\overline{\pi_{sge}} = \pi \cdot \pi_{sge}$, where π is the component of $\overline{\pi_{sge}}$ that is common to all demographic groups. Define $\boldsymbol{\pi}_{sge}$ as the vector of all $\overline{\pi_{sge}}$ parameters, and note that $Y_t(\cdot)$ is homogeneous of degree 1 in $\boldsymbol{\pi}_{sge}$. As a result, the common component π will appear as a Hicks-neutral demand shifter in equation (52).

s , g , e and t - is thus a special case of the model, and the standard Katz and Murphy two schooling group model is a special case that combines the assumption of common demand shifts with the assumption that workers within each schooling group are perfect substitutes in production ($\sigma_g \rightarrow \infty$ and $\sigma_e \rightarrow \infty$). Second, the assumed structure of each schooling group s_t in equations (2) and (3) emphasizes comparisons both within and between gender.

The worker comparisons suggested by the model are clarified by deriving a wage equation that guides my empirical work. The structure of aggregate production in equations (52), (2) and (3), combined with the assumption that the labor market is competitive so that workers' wages equal their marginal product in each year, yields the key wage equation:

$$(4) \quad \ln w_{sget} = \frac{\sigma_e - 1}{\sigma_e} \ln \pi_{sge} + \frac{1}{\sigma_s} \ln Y_t + \frac{\sigma_s - 1}{\sigma_s} \ln A_{st} - \left[\frac{1}{\sigma_s} - \frac{1}{\sigma_g} \right] \ln s_t \\ + \frac{\sigma_g - 1}{\sigma_g} \ln A_{sgt} - \left[\frac{1}{\sigma_g} - \frac{1}{\sigma_e} \right] \ln g_{st} + \frac{\sigma_e - 1}{\sigma_e} \ln A_{sget} - \frac{1}{\sigma_e} \ln g_{set} + \varepsilon_{sget}$$

where $\ln w_{sget}$ is the log wage of a worker in schooling group s , of gender g and in potential experience group e at time t , and where the (empirical) error term ε_{sget} , not explicitly modeled in equations (52), (2) and (3), is included to reflect sampling variation and any other source of variation in log wages not captured in the model.¹⁶ Examination of the RHS of equation (4) reveals that log wages are determined in part by the (time-invariant) relative efficiency of the demographic group (term 1) and an effect that is common to all demographic groups (term 2). If the empirical term ε_{sget} reflects only classical measurement error,¹⁷ the model suggests that all heterogeneity in the time series of log wages is attributable to three sets of supply and demand forces: those common to all workers of the same schooling group (terms 3 and 4), all workers of the same schooling-by-gender group (terms 5 and 6), and all workers of the same schooling-by-gender-by-potential experience group (terms 7 and 8). In the three sections below, I show how equation (4) suggests sequentially-investigating three separate sets of worker comparisons: across potential experience, across gender and finally across schooling.

¹⁶I explicitly include the error term ε_{sget} at this stage so that estimating equations in Sections IV, V and VI can be derived directly from equation (4).

¹⁷This is a standard but strong assumption, and may not be innocuous. For example, a large literature examines the impact of non-market forces, such as the decline in the real value of the minimum wage (Card and DiNardo, 2002; Lee, 1999; Lemieux, 2006b), on the wage structure. In this paper, I abstract from such forces.

I.B.2 Comparing Workers with Different Levels of Potential Experience

Workers of the same schooling and gender have identical values of RHS terms 3-6 in equation (4); all remaining heterogeneity in the time series of log wages, according to the model, is the result of either shifts in supply (captured in the term g_{set}) or shifts in demand (captured by A_{set}) relative to other potential experience groups within each of four schooling-by-gender groups. This observation undergirds Section IV below, in which I use observed changes in relative wages and supplies to estimate changes in relative demands across potential experience groups within each schooling-by-gender group.

The bulk of the literature on the U.S. wage structure, by contrast, either pools male and female workers into the same potential experience category (Acemoglu and Autor, 2011; Autor et al., 2008), or focuses exclusively on male workers when making these comparisons (Card and Lemieux, 2001; Goldin and Margo, 1992; Juhn et al., 1993; Katz and Murphy, 1992; Murphy and Welch, 1992). While the pooling approach allows for the analysis of the entire labor force, it has at least two downsides: it is often theoretically justified by the unattractive assumption that male and female workers are perfect substitutes, and it may degrade the quality of comparisons between potential experience groups, as the Mincerian measure of potential labor market experience may differently-measure actual labor market experience for men versus women.¹⁸ On the other hand, while considering only male workers may sharpen comparisons between potential experience groups, this approach it is often theoretically justified by the undesirable assumption that men and women are not substitutes in production. The current model, in contrast to both approaches, allows men and women to be imperfect substitutes in production and allows for separate comparisons of potential experience groups by gender.¹⁹

Moreover, by relaxing the assumption of common demand shifts, the current model also allows for an analysis of the roles of both supply and demand in determining changes in wages between potential experience groups. This addresses a division in the literature: one set of papers notes that, between 1940 and the late 1980s, wage differentials across both the schooling and experience dimensions tended

¹⁸If attachment to the labor force is generally stronger for men than for women, a unit of potential experience may on average capture fewer units of actual labor market experience among women.

¹⁹The most-similar model in the literature may be Card and Lemieux (2001). Like that paper, the current model allows comparisons of potential experience groups within males: the two male gender aggregates M_{st} from equation (3) are mathematically-equivalent to those defined (for age groups within male workers, rather than potential experience groups) in Card and Lemieux. However, while the assumed structure of aggregate production in equations (52), (2) and (3) mathematically nests the standard model of Katz and Murphy, it does not nest the Card and Lemieux model, which implicitly assumes that male and female workers are not substitutes in production (this would be true if, e.g., aggregate production is Leontief in male and female production).

to move together: shrinking during the "Great Compression" of the U.S. wage structure from 1940 to 1950, then rising together from the early 1960s until the late 1980s, and argues that decreases, then increases, in the returns to labor market "skill," as embodied in both schooling and experience, played a large role in shaping these differentials (Goldin and Margo, 1992; Juhn et al., 1993; Katz and Murphy, 1992; Murphy and Welch, 1992). Recent literature that jointly-examines schooling and experience differentials since the early 1960s has tended to examine the hypothesis that supply shifts alone can explain experience differentials, so that skill-biased technological change is hypothesized to affect only schooling differentials (Autor et al., 2008; Card and Lemieux, 2001). In the empirical sections below, I use the model to examine both supply and demand forces in shaping experience differentials.

Finally, one potential issue with the approach outlined in this section is that a Mincerian potential experience measure - a function of only age and schooling - may mis-measure actual labor market experience among women and, moreover, changes over time in the estimated returns to potential experience among women may reflect not only the returns to actual labor market experience, but shifts in the ratio of actual to potential experience (O'Neill, 2003a; 2003b; O'Neill and Polachek, 2003). In Section IV below, I discuss the potential role of such mis-measurement in explaining my results.

I.B.3 Comparing Female and Male Workers

Workers who have completed the same level of schooling before labor market entry have identical values of RHS terms 3-4 in equation (4); after adjusting for the effects of supply and demand captured in RHS terms 7-8, all remaining heterogeneity in the time series of log wages, according to the model, is the result of either shifts in supply (captured by g_{st}) or demand (captured by A_{sgt}) specific to each gender relative to all workers within the schooling group. In Section V below, I use the supply and demand effects estimated in Section IV, combined with this observation, to estimate changes in the relative demand for female labor within both the college- and high school-educated by comparing observed changes in relative wages and supplies between gender within each schooling group.

This approach is related to a large literature on the evolution of the gender wage gap (Goldin, 1990 provides an excellent overview of the literature), and especially the narrowing of this gap during the 1980s (Blau and Kahn, 1994; 1997; Katz and Murphy, 1992; O'Neill and Polachek, 1993; Smith and Ward, 1989). Overall, this literature suggests that relative demand shifts towards women beginning

in the late 1970s, potentially reflecting, among other sources, industry- and occupation-level shifts in labor demand such as the decline of male-intensive basic manufacturing jobs and the rise of female-intensive sales and clerical work, likely played a large role in the narrowing of the gender wage gap across this period.²⁰ O'Neill and Polachek (1993) suggest that roughly half of the narrowing of the gender wage gap from 1976 to 1990 is due to changes in individual characteristics, while the other half is due to changes in returns to these characteristics. Similarly, decompositions performed by Blau and Kahn (1994; Table 1) imply that changing individual characteristics are responsible for 2/3 of the reduction in the gender wage gap from 1975 to 1987, with the remaining 1/3 attributable to changing returns. Katz and Murphy (1992, Table 6) compute demand shifts derived from employment changes across industry-by-occupation cells across the 1967-1987 period that are roughly 4.3% and 16.1% larger for women than men among the college- and high school-educated, respectively.

While these results suggest the potential importance of shifts in labor demand across gender, and while they could potentially be associated with an assumed structure of aggregate production similar to equations (52), (2) and (3),²¹ the most-common generalizations of the two schooling group framework avoid the comparison of male and female workers, potentially for four reasons. First, while a number of authors suggest that the narrowing of the gender wage gap in the 1980s was driven by shifts in the relative demand for "brains" over "brawn" that are broadly-related to the notion of skill-biased technological change (Beaudry and Lewis, 2014; Galor and Weil, 1996; Weinberg, 2000; Welch, 2000), other authors suggest that gender-specific forces that were largely orthogonal to skill-biased technological change are responsible for this movement in the gender wage differential (Blau and Kahn, 1994; 1997; Card and DiNardo, 2002). Since the standard two schooling group framework is intended, in large part, to examine the effects of skill-biased technological change, its generalizations have tended to abstract from cross-gender comparisons in order to focus on heterogeneity across the schooling dimension. The focus of the current paper, by contrast, is to understand the effects that

²⁰ A number of other sources of labor demand shifts, including decreases in gender-based labor market discrimination, have also been hypothesized, but are more-difficult to identify than these estimated demand shifts, which can be computed from observable changes in industry-by-occupation employment levels. Katz and Murphy (1992) are among the first to suggest the importance of industry and occupational shifts to changes in the gender wage gap during the 1980s.

²¹ Perhaps the closest analogue to the work in this paper is found in section 6 of Blau and Kahn (1997), in which changes in supply and demand for women, relative to men, are computed for workers grouped by skill. In contrast to the approach of this paper, the authors (i) compare genders within three "predicted" skill groups created using a regression methodology, and (ii) rather than estimating an elasticity of substitution and examining implied relative demands, they compute industry-by-occupation (Katz and Murphy) relative demand shifts.

demand shifts within schooling groups - whatever their source - have on the measurement of growth in the relative demand for college-educated workers in the standard model.

A second potential reason is confusion regarding the true structure of aggregate production. A large amount of work supports the hypothesis that, in the post World War II era, the skills gained by attending college are of first-order importance in the labor market (Becker, 1964; Goldin and Katz, 2008; Mincer, 1974). While this appears to imply, as my model assumes, that men and women of the same level of schooling may be appropriately-viewed as direct substitutes in production, some authors suggest that all female labor, including even high-skilled female labor, is better thought of as a substitute with low-skilled male labor (Grant and Hamermesh, 1981; Topel, 1994; 1997). The production assumptions implicit in equation (2) are justified in part, then, by dissenting work that finds strong substitution effects between women and more-skilled men (Acemoglu and Autor, 2004), and finds little impact of shifts in female labor supply on changes in male wage inequality (Blau and Kahn, 1997; Juhn and Kim, 1999).

A third reason that generalizations of the standard model may not directly compare workers of different genders is that variation in the (relative) supply of female labor that is orthogonal to other wage determinants - such as shifts in the relative demand for female labor - and that, therefore, identifies the slope of the relative labor demand curve (and potentially the Allen elasticity of substitution between female and male labor), is difficult to find. The key identification challenge is that changes over time in labor supply are likely more-related to labor force *participation* decisions among women than among men.²² In response to this empirical challenge, rather than attempting to identify a particular value of the elasticity of substitution between female and male labor, in my empirical work I examine a range of plausible values, guided by existing estimates in the literature.²³

Finally, a potential complication in investigating the gender log wage gap is that changes in the (unobserved) relative "quality" of female labor, by cohort, may vary over time. A large amount of work suggests that observed worker characteristics such as education and labor market experience vary across entering cohorts of female labor and are potentially-important factors driving changes in

²²Changes over time in the stock of male labor supply are, conversely, more-likely to be driven by plausibly-exogenous sources such as demography. Goldin (1990, e.g., Figure 5.1) documents the very large increases in female labor force participation since the mid-19th Century.

²³The recent paper by Acemoglu and Autor (2004) uses cross-state variation in female labor supply due to differential mobilization of the U.S. military in World War II. I use this work, as well as that of other authors, to guide the values of the elasticity of substitution between male and female labor I choose to investigate in Section V.

the gender wage gap (Blau and Kahn, 1994; 1997; Goldin, 1990; O'Neill and Polachek, 1993; Smith and Ward, 1989). Similar cross-cohort variation in unobserved worker characteristics presents a well-known empirical challenge in separating changes over time in the *returns* to skills versus changes in the *content* of skills.²⁴ This issue is addressed, in part, in this paper by explicitly modeling schooling and (potential) labor market experience in equations (52), (2) and (3); however, difficult issues remain (Mulligan and Rubenstein, 2008): I discuss the potential influence of unobserved changes in cohort quality on the analysis of this paper in Section V.

I.B.4 Comparing College- and High School-Educated Workers

Although generalizations of the standard model fruitfully permit other comparisons, the model's key use is the comparison of college- and high school-educated workers and, in particular, in delineating the roles of supply and demand in explaining changes in the college wage premium. After estimating the "within schooling group" supply and demand effects associated with RHS terms 5-8 in equation (4), the final step is to compare college- and high school-educated workers. Equation (4) suggests that, after adjusting for RHS terms 5-8, changes in the wage gap between the college- and high school-educated are due either to changes in relative demand (RHS term 3) or supply (RHS term 4). In Section VI below, I examine how the estimated effects captured in RHS terms 5-8 affect the measurement of the relative demand for college-educated labor. While common generalizations of the two schooling group framework imply adjustments for supply effects (e.g., Card and Lemieux, 2001 imply adjusting for a term similar to RHS term 8), the current model requires adjustment for shifts in relative demand within schooling groups, captured in RHS terms 5 and 7. I therefore focus the discussion in this section on how shifts in demand within schooling groups affect the standard model.

Relaxing the assumption of common demand shifts by introducing the parameters A_{sgt} and A_{sget} creates at least two important qualitative departures from the special case of common demand shifts when comparing college- and high school-educated workers: the traditional measures of the relative supply of college equivalent labor (captured by the ratio C_t/H_t) and the relative demand for college equivalent labor (a function of the ratio A_{Ct}/A_{Ht}) both take on new interpretations. In the sections below, I estimate the *empirical* degree to which these measures differ between the common demand shifts special case and the general case in which A_{sgt} and A_{sget} are allowed to freely vary; here, I

²⁴See, for example, the debate in distinguishing changes in the returns to schooling versus change over time in the relative quality of college-educated labor: e.g., Carneiro and Lee (2011) and Juhn et al. (2005).

discuss the *conceptual* differences between the two.

First, consider relative supply. Under the assumption of common demand shifts, the quantity s_t may be thought of as an index that represents the supply of schooling group s equivalent labor, in which the raw labor supplied by each demographic group, captured by each g_{set} , is weighted by a time-invariant efficiency weight π_{sge} . In the general case, raw labor is further weighted by the time-varying weight $A_{sgt} \cdot A_{sget}$, which more-heavily weights the labor supplied by demographic groups within s that are relatively more-demanded during period t . Thus in the general model, the ratio C_t/H_t can be thought of as a relative supply index that, for each schooling group, adjusts effective supply for changes in the returns to gender- and potential experience-specific skills in the labor market.

Second, consider relative demand. Suppose first that the assumption of common demand shifts is not literally true - that is, that A_{sgt} and A_{sget} do not, in reality, equal 1 in every period and for every group, so that the assumption of common demand shifts is properly-viewed as a simplifying abstraction which ignores these shifts. In this case, since the general case explicitly models each A_{sgt} and each A_{sget} , the term A_{st} is then defined in relation to these shifts. That is, in the general case, A_{st} is defined as that component of relative labor demand that is common to all demographic groups within schooling group s after adjusting for demand shifts that are specific to gender (captured by each A_{sgt}) and to gender-by-potential experience (captured by each A_{sget}). Compare this to the case in which the simplifying abstraction of common demand shifts is imposed. In this case, since the terms A_{sgt} and A_{sget} are not explicitly modeled, A_{st} absorbs them implicitly. That is, under the assumption of common demand shifts, A_{st} is not forced to be orthogonal to the shifts in demand that are idiosyncratic to the underlying gender-by-potential experience groups within s ; as a result, it represents both the component of labor demand common to all demographic groups within s and an average of the idiosyncratic components captured in each A_{sgt} and each A_{sget} . In the sections below, I explore the empirical relevance of these points in detail; first, in the next section I lay the foundation for my empirical analysis by defining the data and empirical measures I use in the paper.

1.C Data and Empirical Measures

In this section, I outline the data, basic processing procedures, and the empirical measures of wages, supplies and relative demands that I use in the rest of the paper. My approach is to follow, as closely as possible, the procedures of Autor et al. (2008), which are used in the benchmark work of Goldin and Katz (2008; 2009) and are closely-related to the procedures of Acemoglu and Autor (2011). For

each step below, I note and motivate any changes made to these procedures to fit the current context.

I.C.1 Data

I use the March Current Population Survey (March CPS) data files to measure yearly changes in relative wages and supplies of the U.S. workforce. The March CPS data files, while providing retrospective rather than point-in-time measures of wage earnings, offer the longest yearly data series on wages and labor force participation in the U.S. which, crucially for the current study, provides 18 years of data before the introduction of the IBM PC in 1981 and allows me to construct wage measures that are limited to full-time, full-year (FTFY) workers.²⁵ To further ensure that the basic processing procedures I use are as close as possible to those of the benchmark studies listed above, I begin by using, as my source data files, the cleaned 1964-2009 March CPS data files (which provide information on the earnings years 1963-2008) used in Acemoglu and Autor (2011) and created with procedures very closely-related to those of Autor et al. (2008).²⁶ In Appendix A.1 I describe in detail each of the individual-level variables I use from these source data files.

I use these individual-level data to create empirical measures for 16 "narrow" schooling-by-gender-by-potential experience groups defined by two schooling categories s , two genders g and 4 potential experience categories e . I follow the literature in dividing workers into bins of 0-9, 10-19, 20-29 and 30-39 years of potential labor market experience. In the rest of this subsection, I describe how I aggregate the individual-level source data described here and in Appendix A.1 to create empirical wage, supply and demand measures for each of these 16 narrow demographic groups, as well as each of the broader demographic groups represented in equations (52), (2) and (3).

The empirical measures described below are drawn, as is standard,²⁷ from two separate sub-samples of the March CPS. First, a "wage sample" is limited to FTFY wage and salary workers who possess between 0 and 39 years of potential experience and are between the ages of 16 and 64 during each earnings year. Second a "quantity sample," in contrast, includes all employed workers - not just

²⁵Katz and Murphy (1992) argue that a focus on full-time, full-year workers maximizes the comparability of samples across years, which minimizes the effects of changing sample composition on wage measures. While the Current Population Survey May and Outgoing Rotation Group samples (MORG) (i) provide point-in-time wage measures, and (ii) appear to be less-noisy than those in the March CPS (e.g., Lemieux, 2006b), they begin only in 1973, and do not permit a focus on FTFY workers.

²⁶These files are generously provided by the authors: <http://economics.mit.edu/faculty/dautor/data/acemoglu>.

²⁷Katz and Murphy (1992) were the first to propose using separate samples to measure prices and quantities.

FTFY wage and salary earners - who possess between 0 and 39 years of potential experience. I note the use of each sample in creating the empirical measures below.

I.C.2 Wage Measures

I create wage measures by slightly-modifying the procedure of Autor et al. (2008) to match the demographic groups of interest in this paper. In Appendix A.2, I define the quantity $\widehat{\ln w_{sget}}$, which is intended to capture the log wage of workers in schooling group s , gender g and potential experience group e during earnings year t . While the original wage measures for narrow demographic groups proposed in Katz and Murphy (1992) are simple means taken from the wage sample, the subsequent literature has made two adjustments which I follow in this paper. First, rather than using a raw mean log wage, the log wage measure for each narrow demographic group is adjusted for changes in underlying worker composition using a regression methodology. Second, in order to adjust for changes in composition within college labor (workers with 16 or more years of education) along the schooling margin, the log wage measure for college labor within each of the narrow demographic groups is defined as a fixed-weighted sum of the log wage measures of workers with exactly 16 and 17+ years of education (for a similar approach see, e.g., Goldin and Katz, 2009, Table 2). I fully-describe the implementation of these two steps in the appendix.

Wage earnings for broader demographic groups are commonly-measured as fixed-weighted sums of the wage measures of underlying narrow demographic groups in order to, in effect, hold worker composition constant across the period under study. I follow this procedure by defining the wage measures:

$$(5) \quad \widehat{\ln w_{st}} = \sum_g \left(\widehat{\eta_{sg}} \cdot \widehat{\ln w_{sgt}} \right)$$

$$(6) \quad \widehat{\ln w_{sgt}} = \frac{1}{\widehat{\eta_{sg}}} \cdot \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \widehat{\ln w_{sget}} \right)$$

where I define:

$$(7) \quad \widehat{\eta}_{sg} = \sum_{e=1}^E \widehat{\lambda}_{sge}$$

Here $\widehat{\ln w_{sgt}}$ is intended to capture the log wage of workers in schooling group s and of gender g at time t , while $\widehat{\ln w_{st}}$ measures the log wage of all workers in schooling group s at time t . Each of these broad wage measures is a simple weighted sum of the narrow wage measures $\widehat{\ln w_{sget}}$ using the fixed weights $\widehat{\lambda}_{sge}$. The weight $\widehat{\lambda}_{sge}$ represents the employment, measured as total usual hours worked (as measured in the quantity sample), of workers within schooling group s , gender g and potential experience group e , divided by the employment of all workers within schooling group s , averaged across the entire 1963-2008 sample of earnings years, so that weights sum to 1 within s . The term $\widehat{\eta}_{sg}$ in equation (7) represents, for $g = M$ and $g = F$ respectively, the average employment share of men and women within each schooling group s . It is sometimes convenient to define $\widehat{\eta}_s$ as the share of men within each schooling group s , so that $\widehat{\eta}_{sM} = \widehat{\eta}_s$ and $\widehat{\eta}_{sF} = 1 - \widehat{\eta}_s$ for each s . In Appendix A.2, I outline the formula used to construct each weight $\widehat{\lambda}_{sge}$.

I.C.3 Supply Measures

I follow Card and Lemieux (2001) in constructing empirical supply measures that exactly match the assumed structure of aggregate production.^{28,29} This procedure has two steps. First, I use the quantity sample to create empirical measures of the supply of equivalent labor for each narrow demographic group: for schooling group s , gender g and potential experience e , the empirical measure $\widehat{g_{set}}$ captures the quantity of s equivalent labor supplied by workers of gender g and potential experience e . Recall that for a given gender g and potential experience group e , "equivalent" labor is created by sorting labor supplied by workers of all schooling types into either high school or college equivalent labor. In Appendix A.2, I describe in detail how I construct each $\widehat{g_{set}}$. After constructing these 16 narrow supply measures, I then use equations (2) and (3), in combination with empirical estimates of each

²⁸The literature diverges on whether, when schooling groups are theorized to contain CES aggregates, empirical measures of supply should reflect this CES form (including aggregation using a CES parameter), or whether empirical measures should be simple weighted sums. While recent work (e.g., in the analysis of imperfect substitution between experience groups in Acemoglu and Autor, 2011 and Autor et al., 2008) appears to take the latter approach, here I take the former approach, which more-closely parallels the model.

²⁹The steps described in this sub-section are also similar to Autor et al. (2008) in defining the quantity sample and in creating measures of equivalent labor.

underlying parameter (that I develop below) to create, for each time t , four empirical supply measures \widehat{g}_{st} which denote the effective supply of gender g equivalent labor within schooling group s , and two empirical supply measures \widehat{s}_t which denote the effective supply of schooling group s equivalent labor.

I.C.4 Relative Demand Measures

I construct measures of relative labor demand that are similar to the index of the relative demand for college equivalents developed originally by Katz and Murphy (1992) and used to measure skill-biased technological change from 1940-1996 by Autor et al. (1998). Specifically, I define the index \widehat{D}_t that captures the demand for college equivalent labor relative to high school equivalent labor at time t :

$$(8) \quad \widehat{D}_t = (\widehat{\sigma}_s - 1) \ln \left(\frac{\widehat{A_{Ct}}}{\widehat{A_{Ht}}} \right)$$

, the index \widehat{D}_{st} that captures the demand for female equivalent labor relative to male equivalent labor within schooling group s , at time t :

$$(9) \quad \widehat{D}_{st} = (\widehat{\sigma}_g - 1) \ln \left(\frac{\widehat{A_{sFt}}}{\widehat{A_{sMt}}} \right)$$

and the index \widehat{D}_{sget} that captures the relative demand for equivalent labor of potential experience group e relative to all other potential experience groups within the schooling-by-gender group defined by s and g at time t :

$$(10) \quad \widehat{D}_{sget} = (\widehat{\sigma}_e - 1) \left[\ln \widehat{A_{sget}} - \left(\frac{1}{E - 1} \right) \cdot \sum_{j \neq e} \ln \widehat{A_{sgjt}} \right]$$

\widehat{D}_t is identical to the Katz and Murphy and Autor et al. measures (given the functional form assumptions in equation 52) and is my preferred measure of the relative demand for college-educated labor. The measure depends on an estimate of the elasticity of substitution between college and high school equivalent labor, $\widehat{\sigma}_s$, as well as estimates of the relative demand parameters A_{Ct} and A_{Ht} , which I denote by $\widehat{A_{Ct}}$ and $\widehat{A_{Ht}}$, respectively. In the empirical sections below, I develop estimates of

each of these parameters. The definitions of \widehat{D}_{st} and \widehat{D}_{sget} follow the definition of \widehat{D}_t .

In Appendix A.3, I show that the relative demand measures defined in equations (8), (9) and (10) are each special cases of the same general relative demand measure and do not depend on the simplifying assumption in equations (52), (2) and (3) that technology takes labor-augmenting form (as opposed to share form or a combination of share form and labor-augmenting form). I also show that if the parameter estimates in equations (8), (9) and (10) equal their theoretical counterparts, and if the labor market is competitive, the demand indices defined above can be expressed as:

$$(11) \quad \widehat{D}_t = \sigma_s \ln \left(\frac{w_{Ct}}{w_{Ht}} \right) + \ln \left(\frac{C_t}{H_t} \right)$$

$$(12) \quad \widehat{D}_{st} = \sigma_g \ln \left(\frac{w_{sFt}}{w_{sMt}} \right) + \ln \left(\frac{F_t}{M_t} \right)$$

and

$$(13) \quad \begin{aligned} \widehat{D}_{sget} = & \sigma_e \left[\ln w_{sget} - \left(\frac{1}{E-1} \right) \cdot \ln \sum_{j \neq e} \ln w_{sgjt} \right] \\ & + \ln (g_{set}) - \left(\frac{1}{E-1} \right) \cdot \ln \sum_{j \neq e} \ln (g_{sjt}) \end{aligned}$$

where w_{st} represents the marginal product of schooling aggregate s_t , w_{sgt} represents the marginal product of schooling-by-gender aggregate g_{st} and where w_{sget} represents the marginal product of schooling-by-gender-by-potential experience aggregate g_{set} . That is, each measure has an intuitive interpretation: in each case, relative demand is defined as that component of the relative wage that is unexplained by relative supply.

1.D Changes in the Relative Demand for Potential Experience

In this section, I estimate the shifts in relative demand that, combined with observed shifts in relative supply, are potentially-responsible for changes in the returns to potential experience since the early

1960s. The goal of the section is to understand the shifts in supply and demand between potential experience groups within both college- and high school-educated workers that may affect the measurement of the relative demand for college-educated labor in Section VI. My analysis is guided by the key wage equation of the model, equation (4), which addresses two key issues. First, the model allows for the possibility that shifts in demand (as in Juhn et al., 1993; Katz and Murphy, 1992; Murphy and Welch, 1992), and not solely supply (as in Autor et al., 2008; Card and Lemieux, 2001), may be responsible for changes in the returns to potential experience. Second, the model allows for the possibility that changes in the returns to experience may be different for men than for women (as suggested by Blau and Kahn, 1994; 1997; O’Neill and Polachek, 1993).

In this section, I proceed in three steps: in each successive step, I impose more structure on the data. I begin by observing whether, within four schooling-by-gender groups and across two 21-year periods, wages and supplies moved together or apart. If, for example, the wages and supplies of more-experienced workers within a given schooling-by-gender group both outpaced the leave-out average over a given period, an important role for demand may be implied. Second, I follow the seminal work of Card and Lemieux (2001) in using equation (4) to model heterogeneity in the time series of the college wage premium across more- versus less-experienced workers, and to estimate the key parameter of interest - the elasticity of substitution between potential experience groups, σ_e .³⁰ Before attempting to pin down a particular value of this parameter I first examine, for a wide range of potential values, a set of simple hypothesis tests that indicate whether, for each gender, demand shifts likely play a role in determining cross-potential experience group variation in the college wage premium. I then estimate σ_e for both genders, first by imposing the assumption of common demand shifts, then relaxing this assumption. In the third step, I use my preferred estimate for σ_e from these regressions to estimate the shifts in supply and demand responsible for changes in potential experience wage differentials since the early 1960s. These estimates suggest an important role for shifts in relative demand among all schooling-by-gender groups and, consistent with literature on the narrowing of the gender wage gap in the 1980s, a trend break in relative demand favoring more-experienced workers among women and high school-educated men beginning in the late 1970s. Finally, I briefly show that mis-measurement of actual labor market experience among women appears unlikely to explain these results.

³⁰In their main analysis, Card and Lemieux compare college wage premia among men of different ages. Recent work has tended to modify this approach by examining different potential experience groups, rather than different age groups - a change I maintain in this paper.

TABLE 1
Annual Changes (times 100) in Measures of Log Relative Wages, Supplies and Demands
Between Potential Experience Groups Within and Across Schooling-by-Gender Groups, 1964-2006

A. College Labor						
Gender by Potential Experience Group	1964 - 1985			1985 - 2006		
	Relative Wage	Relative Supply	Relative Demand	Relative Wage	Relative Supply	Relative Demand
Men						
0-9	-0.58	0.42	-2.52	-0.02	-3.05	-3.17
10-19	-0.27	0.73	-0.62	0.37	-2.13	-0.26
20-29	0.32	-0.58	1.05	-0.02	1.85	1.76
30-39	0.53	-0.56	2.09	-0.33	3.33	1.67
Wage-Supply Correlation	-0.89			-0.70		
Gross Wage Change: Mean Demand Share						
	0.69			0.36		
Women						
0-9	-0.09	2.13	1.68	-0.51	-3.94	-6.49
10-19	0.39	3.86	5.81	0.12	-2.34	-1.72
20-29	0.08	-1.73	-1.35	0.26	2.20	3.52
30-39	-0.37	-4.26	-6.14	0.12	4.09	4.69
Wage-Supply Correlation	0.79			0.71		
Gross Wage Change: Mean Demand Share						
	0.52			0.55		
B. High School Labor						
Gender by Potential Experience Group	1964 - 1985			1985 - 2006		
	Relative Wage	Relative Supply	Relative Demand	Relative Wage	Relative Supply	Relative Demand
Men						
0-9	-0.80	2.21	-1.82	0.33	-2.24	-0.56
10-19	-0.09	1.43	0.98	-0.15	-1.23	-1.96
20-29	0.41	-1.76	0.32	-0.19	1.62	0.68
30-39	0.48	-1.89	0.51	0.00	1.85	1.84
Wage-Supply Correlation	-0.93			-0.60		
Gross Wage Change: Mean Demand Share						
	0.31			0.40		
Women						
0-9	-0.57	0.75	-2.12	-0.27	-2.32	-3.68
10-19	0.17	2.39	3.22	-0.31	-1.75	-3.31
20-29	0.24	-1.23	-0.02	0.13	1.57	2.21
30-39	0.17	-1.91	-1.08	0.45	2.50	4.77
Wage-Supply Correlation	-0.29			0.97		
Gross Wage Change: Mean Demand Share						
	0.42			0.63		
C. College, Relative to High School Labor						
Gender by Potential Experience Group	1964 - 1985			1985 - 2006		
	Relative Wage	Relative Supply	Relative Demand	Relative Wage	Relative Supply	Relative Demand
Men						
0-9	0.22	-1.80	-0.70	-0.36	-0.81	-2.61
10-19	-0.18	-0.70	-1.61	0.52	-0.90	1.70
20-29	-0.09	1.17	0.73	0.17	0.23	1.08
30-39	0.05	1.32	1.59	-0.33	1.48	-0.17
Wage-Supply Correlation	-0.38			-0.44		
Gross Wage Change: Mean Demand Share						
	0.48			0.58		
Women						
0-9	0.48	1.39	3.80	-0.24	-1.62	-2.81
10-19	0.22	1.46	2.59	0.43	-0.59	1.59
20-29	-0.17	-0.50	-1.33	0.13	0.63	1.30
30-39	-0.54	-2.35	-5.05	-0.33	1.58	-0.09
Wage-Supply Correlation	0.97			-0.21		
Gross Wage Change: Mean Demand Share						
	0.70			0.52		

I.D.1 Comovement of Relative Wages and Supplies Over 21-Year Periods

In Table 1, I examine movements in relative wages and supplies between potential experience groups both within and across schooling-by-gender groups. To measure relative prices and quantities within schooling-by-gender groups at a given time, I define the deviations:

$$(14) \quad \widehat{rel.wage}_{sget} = \widehat{\ln w_{sget}} - \left(\frac{1}{E-1} \right) \cdot \sum_{j \neq e} \widehat{\ln w_{sgjt}}$$

$$(15) \quad \widehat{rel.supply}_{sget} = \widehat{\ln g_{set}} - \left(\frac{1}{E-1} \right) \cdot \sum_{j \neq e} \widehat{\ln g_{sjt}}$$

and to measure relative prices and quantities across schooling groups, I difference expressions (14) and (15) to define:

$$(16) \quad \widehat{rel.wage}_{get} = \left[\widehat{\ln w_{Cget}} - \widehat{\ln w_{Hget}} \right] - \left(\frac{1}{E-1} \right) \cdot \sum_{j \neq e} \left[\widehat{\ln w_{Cgjt}} - \widehat{\ln w_{Hgjt}} \right]$$

$$(17) \quad \widehat{rel.supply}_{get} = \ln \left(\frac{\widehat{g_{Cet}}}{\widehat{g_{Het}}} \right) - \left(\frac{1}{E-1} \right) \cdot \sum_{j \neq e} \ln \left(\frac{\widehat{g_{Cjt}}}{\widehat{g_{Hjt}}} \right)$$

Equations (14) and (15) are natural measures of relative prices and quantities within each schooling-by-gender group: these expressions capture log differences between a given potential experience group e and the (leave out) average among workers of all other potential experience types with the same level of schooling and gender. Equations (16) and (17) compare the college wage premium and the relative supply of college equivalent labor, respectively, across gender-by-potential experience groups.

In Table 1, I display changes in these measures across two 21-year periods (1964-1985 and 1985-2006).³¹ I examine comovement of relative wages and supplies across each period to shed light on the role of demand shifts. While a negative correlation between relative wages and supplies does not rule out an important role for demand, a positive correlation is instructive. If the structure of

³¹In order to reduce measurement error, for all empirical analysis related to heterogeneity within schooling-by-gender groups, from 1964 to 2003, I use three-year averages of the basic schooling-by-gender-by-potential experience statistics described in the previous sections, and for the year 2006 I use the average of these statistics from the 2005-2008 earnings years. Thus, across the 1963-2008 period of interest, I use data from 15 such averages.

aggregate production is accurately captured by equations (52), (2) and (3) so that equations (14) through (17) represent the "right" comparisons, a positive correlation between relative wages and supplies is strongly-suggestive of important demand effects. A positive correlation may, of course, also indicate model mis-specification (i.e. making the "wrong" comparisons) or the effects of noise. I discuss first male, then female workers.

Men Panels A and B compare changes in the wages of older, more-experienced men to younger, less-experienced men within both the college- and high school-educated. Consistent with literature on increases in the returns to experience during the 1980s, from 1964 to 1985, growth in the wages of more-experienced male workers within both schooling groups significantly outpaced the group average: for example, in both schooling groups, the wages of men with 30-39 years of potential experience grew by approximately 0.5% more annually, or approximately 10% more than the leave-out average over this period.³² From 1985 to 2006 this trend, again within both schooling groups, was reversed: the wages of workers with at least 20 years of experience rose more-slowly, on average,³³ than the wages of less-experienced workers over this period. Relative supply movements among male workers in Panels A and B do not, by themselves, suggest that demand necessarily played a role in these wage changes. Between 1964 and 1985, as the baby boom cohort began to enter the labor market, the relative supply of men with at least 20 years of potential experience declined within both schooling groups, potentially pushing up (own) relative wages. Between 1985 and 2006, as the baby boom cohort matured in the labor market, the supply trend was reversed: the relative supply of workers with 20+ years of potential experience increased, potentially pushing down (own) relative wages. Within each schooling-by-gender group and in each panel of Table 1, the correlation between changes in relative wages and relative supplies across potential experience groups is presented as a useful summary measure that captures whether, on average, relative wages and supplies moved together or apart across each period. The four values of this correlation among male workers in Panels A and B range from -0.93 to -0.60, indicating that, within both college- and high school-educated males, and across both 21-year periods, relative wages and relative supplies covaried negatively.

³²The magnitude of these effects are consistent with, e.g., Katz and Murphy (1992), who report that log wages grew 12% more, on average, among men with 26-35 years of experience relative to men with 1-5 years of experience from 1963 to 1987.

³³Here, and in the rest of this section, I informally refer to simple averages of the changes displayed in the table, as opposed to, e.g., employment-weighted averages.

Panel C compares growth in the college wage premium between less- versus more-experienced men. Consistent with existing evidence (e.g., Panel B of Figure 1 in Katz and Murphy, 1992; Card and Lemieux, 2001) the college wage premium increased more, on average, among groups with fewer than 20 years of potential labor market experience. The sizes of these relative wage gains are similar in scale to those of Panels A and B: for example, the college wage premium among men with fewer than 10 years of potential experience increased roughly 4% more than the average among more-experienced men from 1964 to 1985, while the college premium among men with between 10 and 19 years of potential experience rose roughly 11% more than average from 1985 to 2006. Although differences in the trends of the college wage premium between more- and less-experienced workers are less clear-cut than the wage differences found in Panels A and B, on average the college wage premium rose most-quickly among younger, less-experienced men across the entire 1964-2006 period. Relative supply movements captured in Panel C reveal that, across both periods, the supply of college equivalent relative to high school equivalent labor grew fastest among older men, consistent with the original observation of Card and Lemieux that the slowdown in the educational attainment of the labor force most-strongly affected younger, less-experienced men across these periods. For both periods, the cross-potential experience group correlation between changes in the college wage premium and changes in relative supply among men are negative: -0.38 for the first period and -0.44 for the second.

Panels A, B and C reveal consistent negative comovement between relative wages and relative supplies of different potential experience groups, both within and across schooling groups, across periods of two decades. Taken together, these comparisons are not (themselves) sufficient to imply a role for demand, and are broadly consistent with both the notions that shifts in supply alone, or shifts in both supply and demand, were responsible for changes in relative wages across potential experience groups among men across the 1964-2006 period. Next, I see whether simple comovement of wages and supplies are more-informative about the role of demand within groups of female workers.

Women Inspection of Panels A and B reveals that relative supply shifts within both college and high school equivalent female workers follow the same pattern documented among males: the relative supply of women with at least 20 years of potential experience decreased from 1964 to 1985, then increased from 1985 to 2006. Further, these supply trends are, on average across all schooling groups and time periods, more-dramatic for women and, especially, for college-educated women. For example, on average among college-educated workers, the relative supply of women with 20+ years of experience

decreased by approximately 3% annually over the first period, then increased by approximately 3.1% annually over the second period, compared with a decrease of 0.6% and an increase of 2.6% among males, respectively, over these periods.

Despite the dramatic shifts between more- versus less-experienced female labor in Panels A and B, the corresponding relative wages of women, in general, moved in the *same* direction. The wages of college-educated women with 20+ years of potential labor market experience grew, on average, significantly more-slowly than the wages of less-experienced college-educated women from 1964 to 1985 despite a simultaneous decrease in relative supply, then rose from 1985 to 2006 despite a concurrent relative supply increase. Similarly, from 1985 to 2006, the wages of high school-educated women with at least 20 years of potential experience grew relative to less-experienced women despite growth in relative supply. Panels A and B reveal only one case - that of high school educated women from 1964 to 1985 - in which relative wages and supply moved in opposite directions. In the three other instances just described, the relevant cross-potential experience correlations between relative wages and relative supplies range from 0.71 to 0.97.

Differences in the paths of college wage premia between potential experience groups among women are displayed in Panel C. Between 1964 and 1985, the college wage premium rose significantly faster among women with fewer than 20 years of potential experience. For example, the college wage premium increased 10% more than the leave-out average for women with fewer than 10 years of experience. During this period, the relative supply of college-educated women was also increasing most-quickly among younger, less-experienced women: the correlation between relative wage and supply movements among women across this period is 0.97. From 1985 to 2006, the picture is less-clear: over this period, the college wage premium rose most-quickly among women with between 10 and 39 years of potential experience, and the associated relative wage to relative supply correlation is -0.21. Across the entire 42-year period, I highlight one case, to which I return in detail below: from 1964 to 2006, the college wage premium among women with 10-19 years of potential experience increased roughly 14% more than the average among other potential experience groups within women, despite relative supply concurrently-increasing roughly 18% more than average.

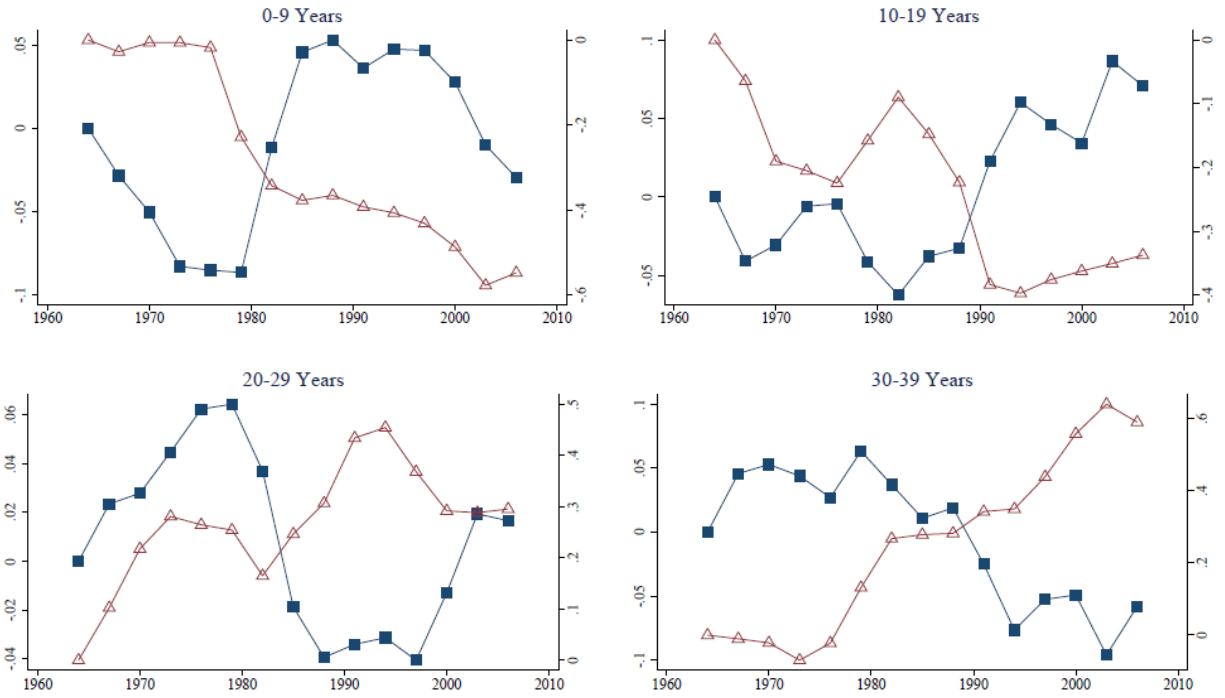
In 3 of 4 cases in Panels A and B, and in 1 of 2 cases in Panel C, relative wages and supplies of potential experience groups among female workers covaried strongly and positively over periods of two decades. If the structure of aggregate production is accurately captured by equations (52), (2) and (3), these results are suggestive that, among female workers, relative demand shifts may be required

to explain observed changes in relative wages between potential experience groups.

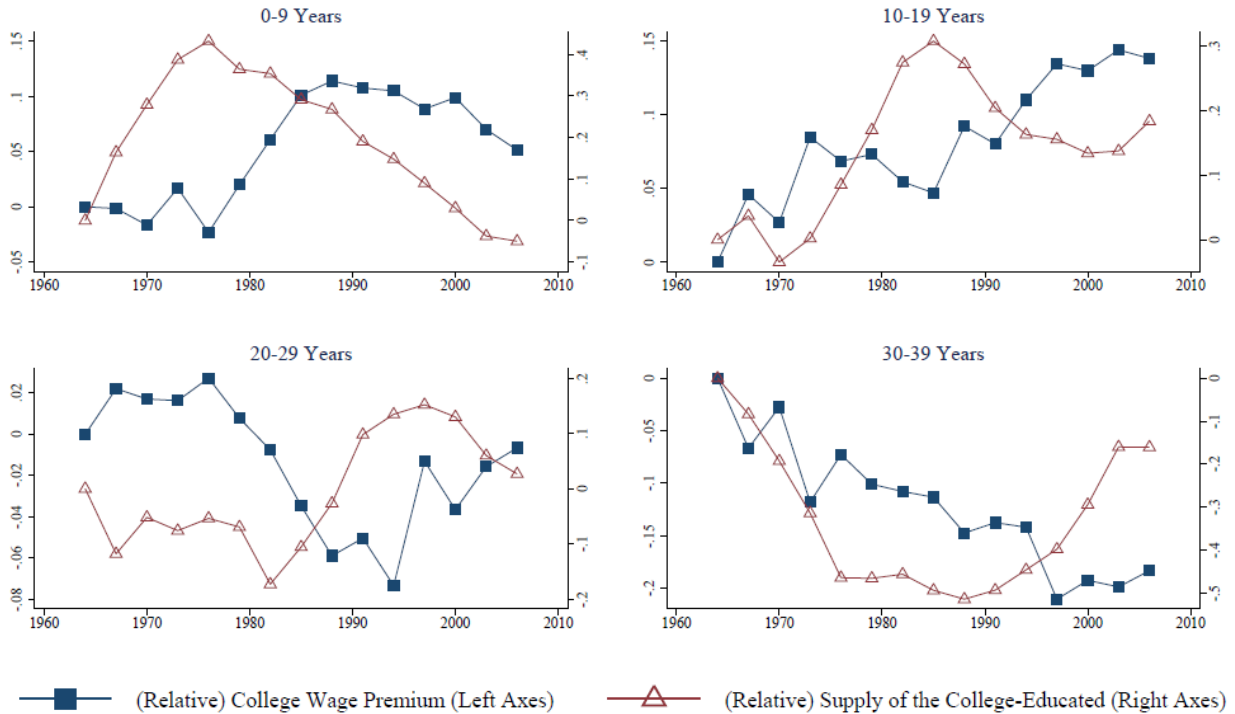
Overall, simple comparisons of 21-year movements in relative wages and supplies are relatively-uninformative about the role of demand shifts in the rise and fall of experience differentials among men, but appear to suggest an important role for demand among women. For example, one explanation of the finding that both wages and supplies rose more-quickly among women with more potential experience from 1985-2006 is that the returns to experience were increasing among women over this period. In order to investigate this hypothesis more-formally, note from equation (3) that, according to the model, workers of different potential experience levels within a given schooling-by-gender group are related by the elasticity of substitution σ_e . Inspection of the key wage equation, equation (4), suggests two steps. First, I must estimate the key the elasticity of substitution between potential experience groups σ_e . Then, with this estimate in hand, I can more-precisely investigate the role of demand shifts within both male and female labor. In the next section, I take the first step: I follow Card and Lemieux (2001) in estimating σ_e by comparing growth in the college wage premium and the relative supply of college-educated labor for different potential experience groups separately within gender.

Figure 1: (Relative) College Wage Premium and Supply of the College-Educated,
for Gender-by-Potential Experience Groups within each Gender

A. Men



B. Women



■ (Relative) College Wage Premium (Left Axes) ▲ (Relative) Supply of the College-Educated (Right Axes)

I.D.2 Estimating a (Card Lemieux) Model of the College Wage Premium

Differencing equation (4) by schooling and potential experience yields an estimating equation that belongs to a class introduced by Card and Lemieux:³⁴

$$(18) \quad rel.wage_{get} = \alpha_{ge} + \frac{\sigma_e - 1}{\sigma_e} \cdot rel.demand_{get} - \frac{1}{\sigma_e} \cdot rel.supply_{get} + \varepsilon_{get}$$

where I have used the notation:

$$\begin{aligned} rel.wage_{get} &= \ln\left(\frac{w_{Cget}}{w_{Hget}}\right) - \left(\frac{1}{E-1}\right) \sum_{j \neq e} \ln\left(\frac{w_{Cgjt}}{w_{Hgjt}}\right) \\ \alpha_{ge} &= \frac{\sigma_e - 1}{\sigma_e} \cdot \left[\ln\left(\frac{\pi_{Cge}}{\pi_{Hge}}\right) - \left(\frac{1}{E-1}\right) \sum_{j \neq e} \ln\left(\frac{\pi_{Cgjt}}{\pi_{Hgjt}}\right) \right] \\ rel.demand_{get} &= \ln\left(\frac{A_{Cget}}{A_{Hget}}\right) - \left(\frac{1}{E-1}\right) \sum_{j \neq e} \ln\left(\frac{A_{Cgjt}}{A_{Hgjt}}\right) \\ rel.supply_{get} &= \ln\left(\frac{g_{Cet}}{g_{Het}}\right) - \left(\frac{1}{E-1}\right) \sum_{j \neq e} \ln\left(\frac{g_{Cjt}}{g_{Hjt}}\right) \\ \varepsilon_{get} &= \left(\frac{\varepsilon_{Cget}}{\varepsilon_{Hget}}\right) - \left(\frac{1}{E-1}\right) \sum_{j \neq e} \left(\frac{\varepsilon_{Cgjt}}{\varepsilon_{Hgjt}}\right) \end{aligned}$$

The relative wage and supply terms from equation (18) are the exact theoretical analogues of the empirical relative wage and supply measures defined in equations (16) and (17) and displayed in Panel C of Table 1. Equation (18) adds structure to the comparisons in Table 1 by noting that, according to the model, relative wages and supplies between potential experience groups are related by a constant elasticity of substitution, σ_e : all else equal, for each gender-by-potential experience group, as relative supply shifts, the resulting joint movements of relative prices and quantities trace out the slope, equal to $-1/\sigma_e$, of an (inverse) relative demand curve associated with equation (18). Further, since the general case of the model allows for demand to vary within schooling-by-gender aggregates, equation

³⁴The closest analogue of equation (18) is Card and Lemieux's "first step" estimating equation (their equation 11).

(18) also makes explicit the notion that demand, captured in the term $rel.demand_{get}$, and not only supply, may play a role in determining relative wage changes.

While Panel C of Table 1 presented long-run changes in the empirical measures defined in equations (16) and (17), equation (18) implies a relationship both in changes and in levels. In Figure 1, I display the full relative wage and supply series defined in equations (16) and (17) separately for each gender-by-potential experience group.³⁵ Under the assumption of common demand shifts, the relative demand term in equation (18) may be safely omitted,³⁶ and all movements in relative wages are assumed to be attributable either to movements in relative supplies or to the term ε_{get} , which captures all sources of variation in relative wages not specified in the model.

Inspection of Figure 1 reveals two interesting patterns that shed light on this hypothesis. First, for *both* male and female workers, over 3- 6- or 9-year periods, movements in relative supplies appear to be associated with opposite concurrent movements in relative wages. I return to this observation below. Second, Figure 1 is consistent with the descriptive statistics in Panel C of Table 1 in that, over 20-, 30- or 40-year periods, while the relative college wage premium and the relative supply of the college-educated are, on average, negatively associated among men, this long-run pattern is much less-clear within women. Recall that these observations suggest that, among female labor, shifts in supply alone may be insufficient to explain changes in potential experience differentials over the 1964-2006 period. In the next section, I use equation (18) to develop a related and simple test in which - separately for each gender-by-potential experience group, and for a broad range of possible values of σ_e within the assumed structure of aggregate production - the hypothesis that supply shifts alone can explain variation in the time series of the college wage premium between potential experience groups can be falsified.

³⁵For ease in making comparisons across time, in Figure 1 I shift each plotted series by a constant so that the value of the series equals zero in the first period. This change is consistent with the form of equation (18), which includes the constant term α_{ge} .

³⁶Recall that in this special case, A_{sget} is assumed to equal 1 for all s , g , e and t , implying that $rel.demand_{get} = 0$ for all g , e and t .

TABLE 2
(Scaled) Covariance Between the Relative Supply and Error Terms of the Standard Equation
for Estimating the Elasticity of Substitution Between Potential Experience Groups σ_e (for Chosen Values of σ_e)

Years of Potential Experience	A. Male Equivalent Workers							Standard Error
	Cobb- Douglas	$\sigma_e=3$	$\sigma_e=4$	$\sigma_e=5$	$\sigma_e=6$	$\sigma_e=7$	Perfect Substitution	
0-9	0.858**	0.191**	0.108*	0.058	0.024	0.001	-0.142**	[0.051]
10-19	0.714**	0.047	-0.036	-0.086	-0.119	-0.143	-0.286**	[0.081]
20-29	0.865**	0.199**	0.115	0.065	0.032	0.008	-0.135	[0.075]
30-39	0.821**	0.154**	0.071**	0.021	-0.013	-0.037	-0.179**	[0.023]
Years of Potential Experience	B. Female Equivalent Workers							Standard Error
	Cobb- Douglas	$\sigma_e=3$	$\sigma_e=4$	$\sigma_e=5$	$\sigma_e=6$	$\sigma_e=7$	Perfect Substitution	
0-9	0.912**	0.246**	0.162*	0.112	0.079	0.055	-0.088	[0.069]
10-19	1.130**	0.463**	0.380**	0.330**	0.297**	0.273*	0.130	[0.113]
20-29	0.833**	0.167*	0.083	0.033	0.000	-0.024	-0.167*	[0.070]
30-39	1.108**	0.442**	0.358**	0.308**	0.275**	0.251*	0.108	[0.105]

Can Supply Shifts Alone Explain Variation in the College Wage Premium? Equation (18)

can be re-written as:

$$(19) \quad \eta_{get} = \frac{\sigma_e - 1}{\sigma_e} \cdot rel.demand_{get} + \varepsilon_{get} = -\alpha_{ge} + rel.wage_{get} + \frac{1}{\sigma_e} \cdot rel.supply_{get}$$

where the term η_{get} represents the error term of a "short" or "standard" variant of equation (18) that omits any terms intended to capture relative demand, so that $rel.demand_{get}$ is grouped together with the unobserved error term ε_{get} . If relative supply shifts alone are sufficient to explain all movements in relative wages - both short run (3-, 6- and 9-year) and longer run (20-, 30- and 40-year) - the short variant of equation (18), with the associated error term η_{get} , may be appropriate to estimate σ_e by comparing relative wages to relative supplies: under this null hypothesis, the correlation between relative supply $rel.supply_{get}$ and the error term η_{get} is zero.³⁷ If, however, factors other than relative supply affect relative wage movements, the correlation between relative supply and the error term of the short regression may be non-zero.

Expression (19) suggests a simple test of the null hypothesis. A value of σ_e , combined with the observed empirical analogues of $rel.wage_{get}$ and $rel.supply_{get}$, define a unique time series of the error term η_{get} .³⁸ Thus, for a given gender-by-potential experience group, and for a chosen value of σ_e , the null hypothesis that the correlation between $rel.supply_{get}$ and the error term η_{get} equals zero can be directly tested. I take $\sigma_e = 4.926$ from Card and Lemieux as the benchmark estimate from the literature;³⁹ in Table 2 I display, separately for 8 gender-by-potential experience groups, a measure of the desired covariance, $Cov(\eta_{get}, rel.supply_{get})$ scaled by $Var(rel.supply_{get})$, for a set of chosen values for σ_e that range from the Cobb-Douglas case ($\sigma_e \rightarrow 1$) to the case of perfect substitutability

³⁷I discuss in detail the assumptions required to estimate σ_e using equation (18) in the next section.

³⁸Formally, without knowledge of the relative efficiency gap α_{ge} , an estimate of a unique *set* of possible time series of η_{get} is implied. Because I am interested in the time-varying component of η_{get} , in practice I shift each value of the time series of η_{get} so that the series equals 0 in the first period.

³⁹I choose the estimate from column 1 of Table 3 of Card and Lemieux (2001) as my benchmark value because (i) in limiting their main analyses to men, they examine sub-aggregates identical to M_{st} from equation (3) and thus estimate a parameter that, in the case of no differential demand shifts within schooling groups and replacing potential experience groups with age groups, is identical to σ_e and (ii) the column 1, Table 3 estimate is from their "first step" procedure which is similar in form to my equation (18). All estimates of σ_e from Card and Lemieux are close to 5.

$(\sigma_e \rightarrow \infty)$.⁴⁰

Inspection of Table 2 suggests a discrepancy between male and female equivalent labor. Among male equivalent workers (Panel A), for no potential experience group can the null hypothesis of a zero covariance be rejected at the 5% level if the true value of σ_e is 5, the "best guess" estimate from the literature. This is not surprising: the Card and Lemieux estimates of σ_e are produced, among a sample of male workers, by fitting equations similar to the short version of equation (18) - a process that mechanically creates a zero correlation. Of the 20 hypothesis tests among male equivalent labor associated with values of σ_e between the values of 3 and 7, only 5 can be rejected at the 5% level among male equivalent workers. Among female equivalent workers (Panel B), by contrast, 13 of 20 such hypothesis tests can be rejected at the 5% level. Two groups - women with 10-19 and 30-39 years of potential labor market experience - show substantial and positive correlation between relative supply and the error term of the short regression for each of these cases, including the benchmark case of $\sigma_e = 5$. Moreover, the observed covariance is positive for 19 of 20 hypothesis tests associated with values of σ_e between the values of 3 and 7 among women. Overall, Table 2 suggests that if equations (52), (2) and (3) accurately-capture the true structure of aggregate production, and if the true value of σ_e is approximately 5, then: (i) the short variant of equation (18) may be appropriate to analyze the prices and quantities of labor supplied by men, but is mis-specified for women, and (ii) wage determinants that are unaccounted for in the short regression are positively-correlated with changes in relative supply.

Equation (19) suggests two types of explanations for the non-zero correlations observed among female workers in Panel B of Table 2. First, sources of variation in the college wage premium outside of the model, captured in the empirical term ε_{get} , may be positively-correlated with relative supply. Second, unobserved shifts in demand, captured by the term $rel.demand_{get}$, may also explain the results. Although it is impossible to rule out explanations of the first sort, the strong and consistently-positive sign of correlations between relative wages and supplies among women in Tables 1 and 2 are consistent with the second explanation.⁴¹ In the next section, I further examine this second hypothesis.

⁴⁰I divide my measure of covariance by $Var(rel.supply_{get})$ in order to match the form of relevant regression coefficients derived from equation (18). I calculate the scaled covariance measure using a bivariate regression of the empirical analogue of η_{get} onto the empirical analogue of $rel.supply_{get}$, and I use the robust standard error of the bivariate regression coefficient to define its confidence interval. As a result, the confidence interval associated with a given gender-by-potential experience group is mechanically identical for each value of σ_e .

⁴¹For example, even in the extreme case of perfect substitution in Panel B of Table 2 - which essentially ignores the

Modeling Shifts in Relative Demand for the College-Educated To motivate the empirical approach of this section, consider the time path of relative wages and relative supplies among female workers with between 10 and 19 years of potential experience in the top right plot of Panel B in Figure 1. For this group, relative wages and relative supplies have grown together over periods of 20, 30 or 40 years which, as noted above, suggests a potential role for relative demand across these intervals. Additionally, over shorter periods of 3, 6, and 9 years, relative wages appear to covary negatively with relative supplies for this group. One potential explanation for these observed patterns is that demand for the college-educated among women with 10-19 years of potential experience has steadily grown relative to other groups over longer periods, explaining the longer-run positive comovement, while fluctuations in relative supplies over shorter periods drive simultaneous and opposite fluctuations in relative wages, explaining the shorter-run negative comovement.

I incorporate the possibility of a smooth long-run trend in relative demand into equation (18) using an approach that mirrors the original logic of Katz and Murphy in modeling long-run trends in the relative demand for college workers. I assume a specific functional form for the time series of relative demand within each gender-by-potential experience group:

$$(20) \quad rel.demand_{get} = f_{ge}(t)$$

where $f_{ge}(\cdot)$ is a gender-by-potential experience group-specific polynomial that equals 0 under the assumption of common shifts and which I alternately assume is of degree 1 or 2 to capture linear or quadratic relative demand trends across the 1964-2006 earnings years, respectively. While linear and quadratic trends may be able to capture shifts in relative demand over periods of 20 years or longer, they will not capture any short-run demand movements. Substituting for $rel.demand_{get}$ in estimating equation (18) using expression (20) then suggests that the elasticity of substitution σ_e may be recovered from the coefficient associated with relative supply: intuitively, if changes in relative demand over 20- to 40-year periods are captured by expression (20), then if the remaining fluctuations in relative wages over 3- to 9-year periods reflect shifting supply, the movements of relative wages and relative supplies will trace the slope of the (inverse) relative demand curve, $-1/\sigma_e$.

effects of relative supply on relative wages, only 1 of the 4 hypothesis tests (for female workers with between 20 and 29 years of potential experience) suggests a negative and significant correlation between relative supply and the error term.

The key empirical assumption associated with this approach is that, after accounting for long-run shifts in relative demand, short-run fluctuations in relative supply are uncorrelated with other potential drivers of short-run relative movement in the (relative) college wage premium, including short-run fluctuations in relative demand. This assumption will hold if, e.g., short run movements in the relative supply of college and high school equivalent workers are driven mostly by educational investment decisions made by workers before labor market entry and are therefore likely to be insensitive to short-run changes in the labor market. It will not hold if, e.g., these movements reflect differential labor force participation decisions, which may be related to relative demand shifts.⁴²

To estimate σ_e separately for each gender-by-potential experience group, equation (20) can be directly-substituted into equation (18) and implemented in one step. In order to estimate σ_e in regressions that pool observations across these groups, I use a two-step procedure. First, I regress the empirical analogues of the relative wage and relative supply terms in equation (18) on the polynomial in equation (20). I then regress the wage residuals from this first step regression on the supply residuals and on dummies for each gender-by-potential experience group. I show in Appendix A.5 that, if changes in relative demand are accurately-captured by equation (20), this procedure results in consistent estimation of the slope of the (inverse) relative demand curve in equation (18). I report all estimates of this slope - separate and pooled (two-step) - in Table 3.

I discuss the estimation of σ_e in four steps. First, I first discuss estimates, among both male and female workers, recovered from "standard" or "short" variants of equation (18) that are suggested by the assumption of common demand shifts. Second, I examine how these estimates change among female workers when relative demand is assumed to be approximated by a quadratic polynomial. Third, I discuss the economic intuition behind differences in these two estimates. Finally, I present my preferred estimate for σ_e .

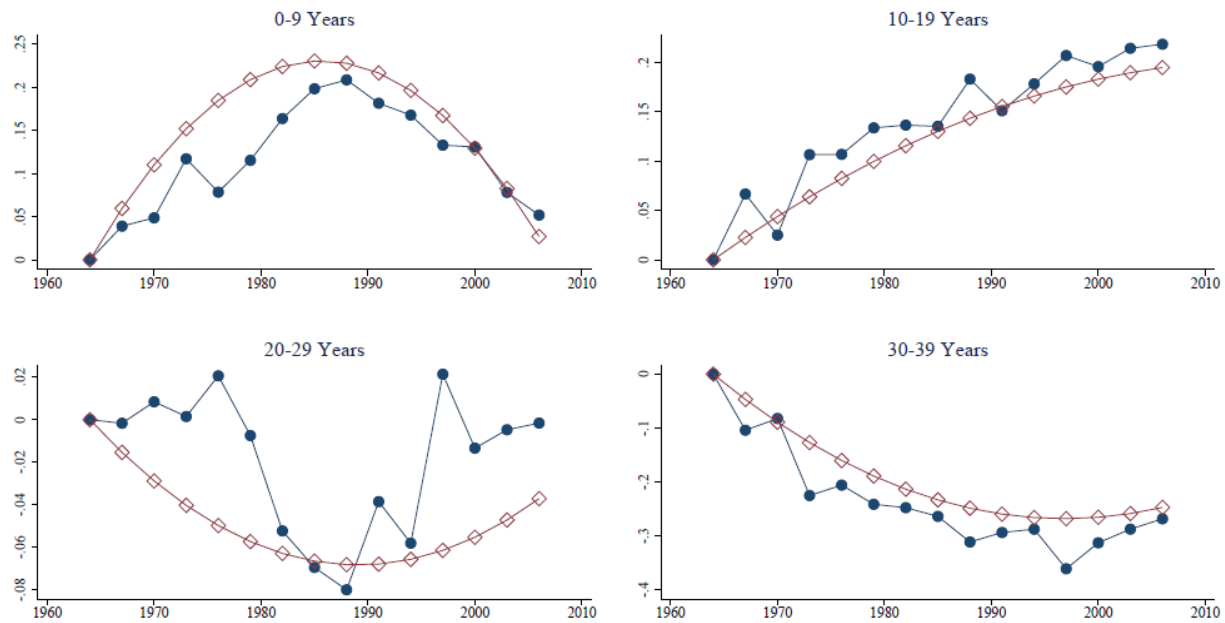
⁴²When defined with respect to the entire U.S. labor force (as opposed to with respect to gender-by-potential experience groups, as in equation 18 and, e.g., Card and Lemieux, 2001) this assumption is supported by instrumental variables evidence (Ciccone and Peri, 2005; Heckman et al., 1998).

TABLE 3
Estimating a Model of Changes in the College Wage Premium: Comparing Potential Experience Groups within Each Gender

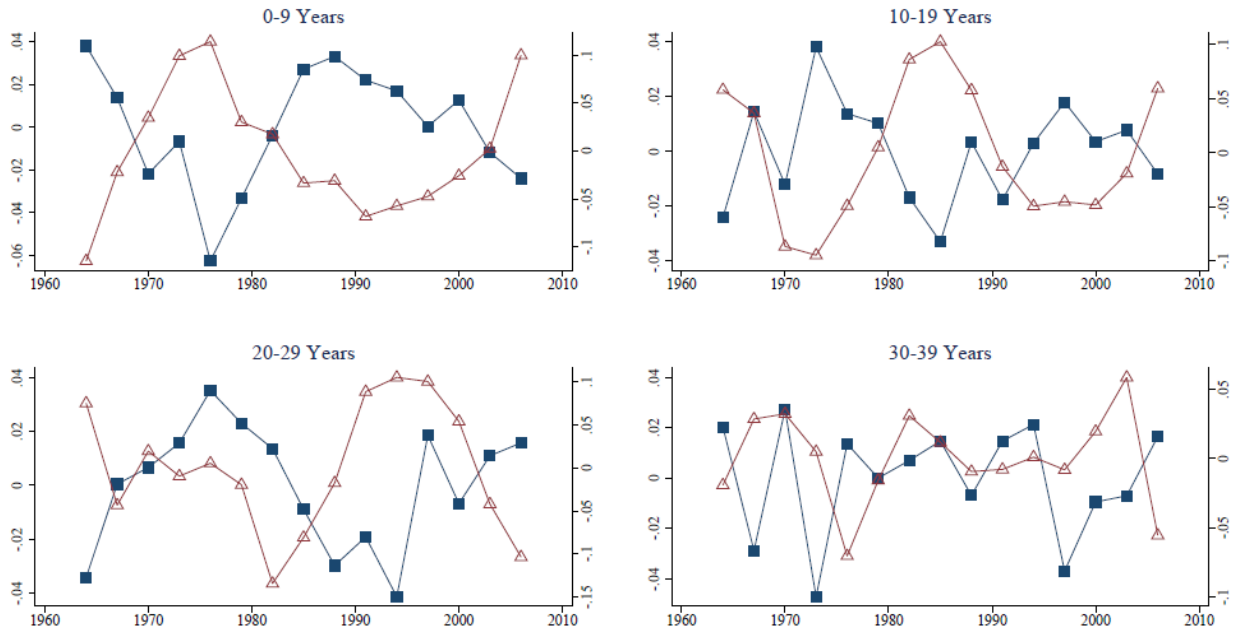
A. Men														
A.1 "Standard" (No Demand Trend)					A.2 Linear Demand Trend					A.3 Quadratic Demand Trend				
	By Years of Potential Experience				By Years of Potential Experience				Pooled, Two-Step Estimate	By Years of Potential Experience				Pooled, Two-Step Estimate
	0-9	10-19	20-29	30-39	0-9	10-19	20-29	30-39		0-9	10-19	20-29	30-39	
Relative Supply	-0.142* [0.051]	-0.286** [0.081]	-0.135 [0.075]	-0.179** [0.023]	-0.175** [0.024]	-0.375* [0.133]	-0.209 [0.116]	-0.075 [0.103]	-0.131 [0.135]	-0.363* [0.137]	-0.266** [0.058]	-0.073 [0.113]	-0.041 [0.107]	-0.189** [0.055]
Time						-0.004 [0.002]	0.001 [0.001]	-0.001 [0.001]	-0.001 [0.003]	-0.003 [0.004]	-0.006** [0.002]	-0.001 [0.004]	0.002 [0.003]	
Time ² / 100										-0.002 [0.008]	0.015** [0.003]	0.000 [0.007]	-0.010 [0.006]	
Constant	0.048* [0.016]	0.088** [0.013]	0.004 [0.009]	-0.128** [0.010]	0.000 [0.011]	0.178* [0.081]	0.057 [0.031]	0.022 [0.022]	-0.093 [0.104]	0.000 [0.101]	0.167 [0.086]	0.022 [0.034]	-0.057 [0.079]	0.000 [0.010]
Observations	15	15	15	15	60	15	15	15	15	15	15	15	15	60
R-Squared	0.355	0.619	0.187	0.698	0.469	0.449	0.638	0.229	0.702	0.453	0.892	0.229	0.773	0.143
R-Squared: Estimated vs. Implied ($\alpha_5=5.025$) Relative Demand Series														
B. Women														
B.1 "Standard" (No Demand Trend)					B.2 Linear Demand Trend					B.3 Quadratic Demand Trend				
	By Years of Potential Experience				By Years of Potential Experience				Pooled, Two-Step Estimate	By Years of Potential Experience				Pooled, Two-Step Estimate
	0-9	10-19	20-29	30-39	0-9	10-19	20-29	30-39		0-9	10-19	20-29	30-39	
Relative Supply	-0.088 [0.069]	0.130 [0.113]	-0.167* [0.070]	0.108 [0.105]	0.008 [0.057]	0.042 [0.067]	-0.111 [0.054]	-0.088 [0.088]	0.031 [0.032]	-0.347** [0.080]	-0.183* [0.079]	-0.129 [0.076]	-0.151 [0.112]	-0.209** [0.043]
Time						0.003** [0.001]	0.003** [0.000]	-0.001 [0.001]	-0.004** [0.001]	0.017** [0.002]	0.006** [0.002]	-0.004* [0.002]	-0.013* [0.005]	
Time ² / 100										-0.039** [0.006]	-0.006 [0.004]	0.009* [0.004]	0.020 [0.012]	
Constant	0.055 [0.033]	-0.003 [0.021]	-0.022 [0.011]	0.020 [0.061]	-0.000 [0.013]	-0.067 [0.032]	-0.036* [0.013]	0.008 [0.022]	0.070* [0.023]	0.057* [0.025]	-0.047** [0.014]	0.022 [0.024]	0.047 [0.030]	0.000 [0.005]
Observations	15	15	15	15	60	15	15	15	15	15	15	15	15	60
R-Squared	0.081	0.099	0.295	0.088	0.000	0.549	0.863	0.385	0.863	0.906	0.887	0.563	0.874	0.309
R-Squared: Estimated vs. Implied ($\alpha_5=5.025$) Relative Demand Series														
C. Men and Women: Quadratic Demand Trend, Pooled Two-Step Estimate														
Regressor														
(31)														
Relative Supply														
-0.199** [0.035]														
Constant														
0.000 [0.005]														
Observations														
120														
R-Squared														
0.199														

Figure 2: Demands and Residuals Within Female Workers, by Potential Experience Group

A. (Relative) Demand for College-Educated Workers: Implied and Estimated Quadratic



B. (Relative) Wage and Supply Residuals, Adjusted for Quadratic Time Trend



Estimates of σ_e Under the Assumption of Common Demand Shifts Panels A.1 and B.1 of Table 3 display estimates of the slope of the (inverse) relative demand curve from equation (18) that, as is standard, do not account for the term $rel.demand_{get}$. Among male workers, separately-estimated point estimates presented in columns 1 through 4 range from -0.286 to -0.142, and the pooled estimate in column 5 yields a point estimate of -0.175, which implies an estimated elasticity of substitution of 5.714. These estimates are broadly similar to estimates of σ_e from samples of male workers from Card and Lemieux, all of which are near 5, including their "first step" estimate of 4.926.⁴³ Moreover, among males, the fit of the short version of equation (18) is quite good: the R-squared statistic associated with the second step regression in column 5 is 0.496, implying that relative supply shifts play an important role in shaping relative wage changes between potential experience groups among male workers.⁴⁴ Among females however, slope estimates are impossible to take literally within the model: separately-estimated coefficients in columns 16 through 19 range from -0.167 to 0.130, with an associated pooled estimate of 0.008. Given the assumed structure of production, an upward-sloping relative demand curve associated with regression (18) is uninterpretable within a framework in which the CES elasticity σ_e must be (strongly) positive. Unsurprisingly, the R-squared estimate reported in column 20 is close to zero.

Estimates of σ_e Among Women: Relaxing the Common Demand Assumption In Panel B.3 of Table 3, by contrast, I relax the assumption of common demand shifts among female workers by assuming that $rel.demand_{get}$ for each potential experience group follows a quadratic time trend across the 1964 to 2006 earnings years. Point estimates of the slope of the inverse relative demand curve associated with equation (18) in columns 26 through 29 range from -0.347 to -0.129, and the pooled, two-step estimate of -0.209 implies an elasticity of substitution of 4.785. On average, across columns 26 through 29, estimates of the coefficients of the quadratic function $[(\sigma_e - 1) / \sigma_e] \cdot f_{ge}(t)$ are precisely-estimated, and 4 of these 6 coefficients are different from 0 at a 5% significance level, potentially

⁴³Card and Lemieux examine a shorter time period (1970 to 1995) and differently-defined demographic subaggregates (age groups, rather than potential experience groups) in estimating σ_e .

⁴⁴Note that because, in my empirical approach, (i) I examine relative wages and supplies in equation (18) by differencing, rather than using time fixed effects, and (ii) the second-step regression features residualized variables on both the LHS and RHS, the R-squared statistic is a "purer" measure of fit than the R-squared statistics reported, e.g., in Column 1 of Table 3 of Card and Lemieux, whose reported R-squared of 0.97 among males for the analogous model implicitly includes the effects of time fixed effects and constants. My approach also has the advantage that the reported R-squared statistics from all second-step regressions are comparable despite differences in demand assumptions across specifications.

indicating that relative demand across the 1964 to 2006 earnings years may be approximated by a quadratic function of time.

Panel A of Figure 2 further investigates this hypothesis by plotting, separately for each potential experience group within females, the estimated quadratic function of time as recovered from the coefficient estimates in columns 26 through 29 against the implied relative demand series recoverable from equation (18), given the value $\sigma_e = 5.025$.⁴⁵ Inspection of these plots reveals that, for women with 0-9, 10-19 and 30-39 years of potential experience, the implied series of $rel.demand_{get}$ is well-captured by a smooth quadratic trend. In columns 26 through 29, I display a simple measure of fit between each estimated and implied relative demand series: the R-squared statistic from a regression of implied on estimated relative demands. For these three potential experience groups, this statistic ranges from 0.791 to 0.926, implying that shifts in $rel.demand_{get}$ between these potential experience groups over the 42 year period under consideration are characterized by smooth changes, rather than short-run fluctuations. For women with 20-29 years of potential experience, the associated R-squared statistic describing the fit between estimated and implied relative demands is a much lower 0.322; the bottom left plot in Panel A of Figure 2 reveals that, for this group, implied relative demand for college graduate labor sharply declined between 1976 and 1982, and sharply rose from 1988 to 1997: these sharp changes are not consistent with smooth changes in relative demand for this group.

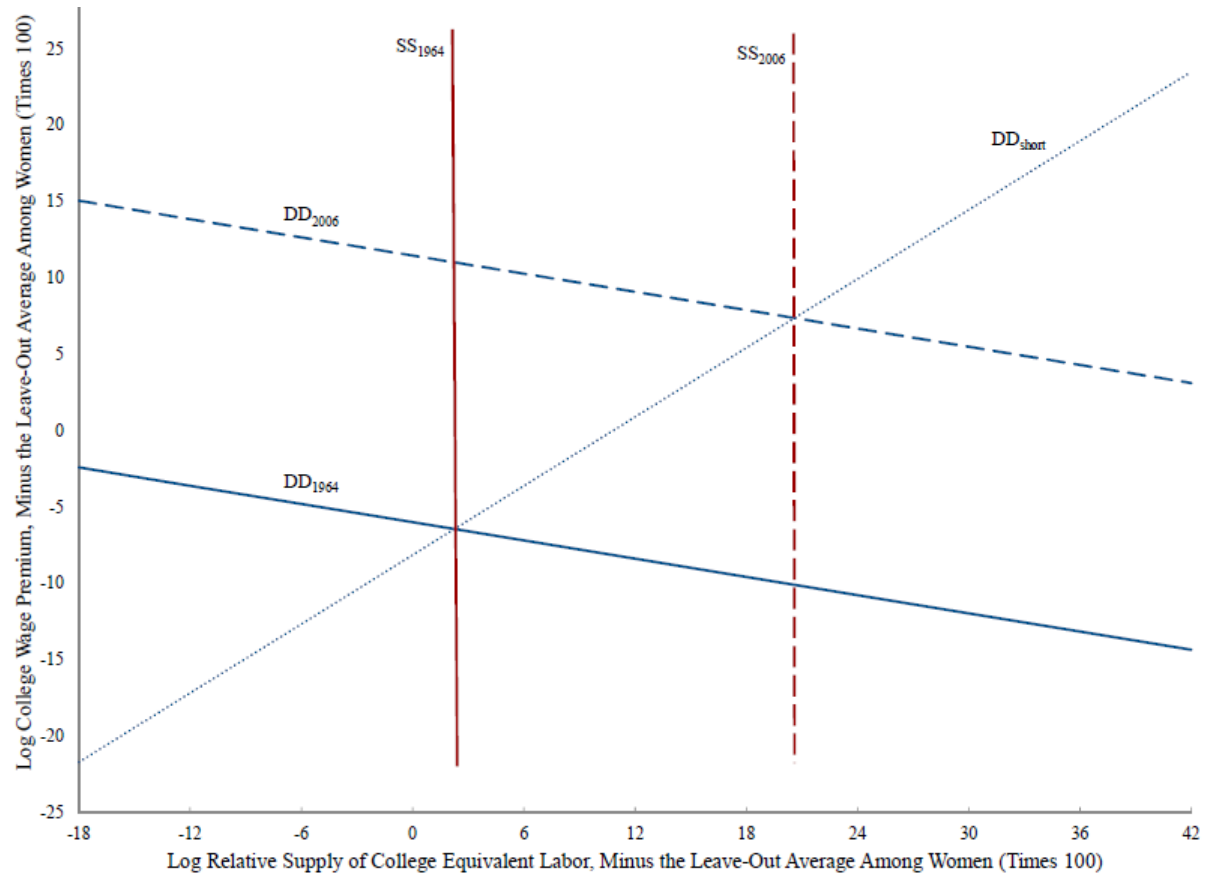
Panel B of Figure 2 examines the relationship between relative wages and supplies after accounting for smooth relative demand changes: it plots, for each potential experience group within females, the wage and supply residuals created from first step regressions in which wages and supplies are regressed on a quadratic time trend. For all four potential experience groups, these plots confirm the suggestive evidence from Panel B of Figure 1 that, after accounting for long-run changes in relative demands, fluctuations in relative wages and supplies over 3-, 6- and 9-year periods are clearly and negatively associated.

Finally, the pooled two-step estimates presented in column 30 are associated with an R-squared statistic of 0.309 which is comparable to (though lower than) the R-squared statistic of 0.469 presented in column 5. This suggests that, after adjusting for longer-run shifts in relative demand, short-run shifts in relative supplies play an important role in determining short-run changes in relative wages

⁴⁵In creating both the estimated and implied relative demand functions, I use my preferred estimate of σ_e , 5.025, derived from column 31 of Table 2, because of its relative precision. For each potential experience group, I shift each value in a given series by a constant so that the series equals zero in the first period.

between potential experience groups among women.

Figure 3: (Relative) Supply and Demand for College-Educated Labor:
Women with 10-19 Years of Potential Labor Market Experience



Does the Assumption of Common Demand Shifts Bias Estimates of σ_e ? First row coefficient estimates in Table 2 are intended to capture the slope of the inverse relative demand curve defined in equation (18). The results discussed so far show that imposing the assumption of common demand shifts among women results in estimates that suggest an inverse relative demand curve that slopes upwards (column 5), while relaxing this assumption using a quadratic version of equation (20) results in a standard downward-sloping curve (column 30). Equation (18) suggests a simple economic interpretation of this measurement gap. If the simplifying assumption of common demand shifts is imposed, but is incorrect, so that the true value of $A_{sget} \neq 1$ for some s , g , e and t , and if no term is included in the estimation of equation (18) to capture variation in demand, the term $rel.demand_{get}$ will appear as part of the error term as an omitted variable. In Appendix A.4, I show that if ε_{get} is uncorrelated with relative supply,⁴⁶ omission of $rel.demand_{get}$ yields the bias:

$$(21) \quad Bias \left[\widehat{\sigma}_e^{short} \right] = -\frac{\sigma_e}{\sigma_e - 1} \cdot \left[\frac{Cov \left(rel.demand_{get}, rel.supply_{get} \right)}{Var \left(rel.supply_{get} \right)} \right]^{-1}$$

where $\widehat{\sigma}_e^{short}$ represents the estimate of σ_e implied from the short variant of equation (18) in which $rel.demand_{get}$ is omitted. Equation (21) makes clear that, if long-run changes in relative demand and relative supply positively-covary, and if the true value of σ_e is greater than 1,⁴⁷ incorrectly assuming common demand shifts (which, in turn, suggests the standard short regression) yields an estimate of σ_e that is biased downwards, and may potentially have a negative value.

Figure 3 provides economic intuition behind equation (21). The intersection of the supply and demand curves in the bottom left of the figure (SS₁₉₆₄, DD₁₉₆₄ and DD_{short}) document that in 1964 the college wage premium and the relative supply of college equivalent labor among women with 10-19 years of potential experience was roughly 2.3% higher and 6.4% lower, respectively, than the leave-out average among women. Across the next 42 years, as documented above, the relative supply of college-educated workers among this group increased roughly 18% more than average, yet despite this supply increase, the relative wage concurrently increased by roughly 14%. The intersection of the supply and demand curves in the top right of the figure (SS₂₀₀₆, DD₂₀₀₆ and DD_{short}) document this

⁴⁶This assumption may not be harmless: bias in σ_e is potentially-attributable to a wide range of sources; in this section, I consider one potential source: omission of relevant shifts in relative demand.

⁴⁷Results in this paper and estimates from the literature suggest that this is the empirically-relevant case.

joint increase. Figure 3 shows that, if the relative demand curve is allowed to shift (from DD_{1964} to DD_{2006}), this long-run positive comovement of relative wages and supplies is fully-consistent with a downward-sloping relative demand curve.⁴⁸ However, if relative demand is forced to remain fixed, the only explanation for the two observed relative wage and supply observations plotted in the figure is an upward-sloping relative demand curve, which I plot as DD_{short} in Figure 3.

A Preferred Estimate for σ_e Comparison of columns 5, 10 and 15 reveals that accounting for long-run shifts in relative demand does not, in these pooled regressions, change the estimated slope of the inverse relative demand curve among men. Slope estimates in these columns range from -0.175 to -0.189, implying estimates of σ_e that range from 5.291 to 5.714. However, 7 of 8 R-squared statistics that describe the fit between estimated and implied relative demands in columns 6 to 9 and 11 to 14 fall below 0.18, suggesting that the time series of relative demand among men may not be accurately-captured by smooth time trends. Similarly, these same statistics from columns 21 and 23 reveal that the assumption of a linear relative demand trend among women does a poor job of capturing shifts in relative demand for women of 0-9 and 20-29 years of potential experience: a point that is reinforced by the distinctly non-linear implied relative demand series plotted in Panel A of Figure 2.

Taken together, estimates in Panel A confirm the original findings of Card and Lemieux that relative wage changes between potential experience groups among males are well-captured by shifts in relative supply within a CES aggregate characterized by an elasticity of substitution of approximately 5. These estimates are also fully-consistent with an important role for relative demand - a point I address in the next section. Estimates in Panel B suggest that relative wage changes between potential experience groups among females are well-captured by smooth changes in demand over 20- and 40-year periods, combined with fluctuations in relative supply over 3- to 9-year periods, within a CES aggregate that is also associated with an elasticity of substitution of approximately 5. In Panel C, I combine male and female labor, allowing for a quadratic time trend to capture relative demand changes within each gender-by-potential experience group. The estimated slope of -0.199 reported in column 31 is associated with my preferred estimate for σ_e of 5.025, which I denote by $\widehat{\sigma_e}$ and is very close to the estimates of Card and Lemieux. In the next section, I use this estimate, together with the model, to infer shifts in relative demand between potential experience groups within each schooling-by-gender

⁴⁸In the figure, I display the relative demand curves associated with my preferred estimate of σ_e of 5.025, derived from column 31 of Table 2. This value is associated with an inverse relative demand curve slope of -0.199.

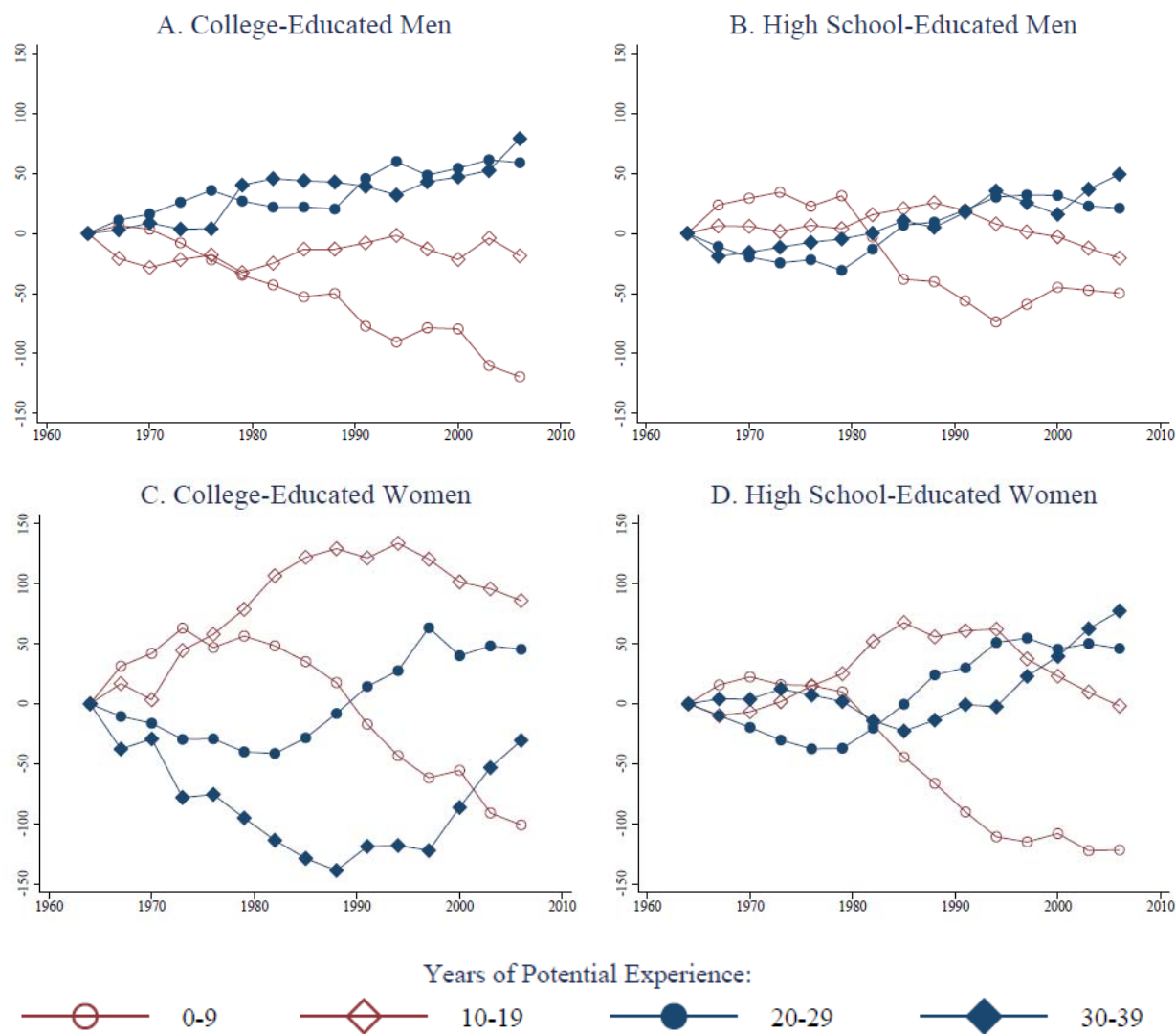
group, and to understand the causes of changes in potential experience differentials.

I.D.3 Estimated Changes in the Relative Demand for Potential Experience

Comparisons of observed relative wages and supplies carried out so far - either completely reduced form (Table 1 and Figure 1), flexibly-applying the model (Table 2), or used to estimate a specific value of the parameter σ_e (Table 3) have suggested that relative demand may play an important role in determining changes in experience differentials among women. By contrast, less light has been shed on the role of demand shifts among male workers. In this section, I fully-apply the model: I show that my preferred estimate $\widehat{\sigma_e}$, combined with the key wage equation, equation (4), yields estimates of shifts in the relative demand for potential experience within each schooling-by-gender group, and specific inferences about the causes of changes in the return to potential experience.

I organize this section into three parts. First, I show how shifts in relative demand for potential experience are estimated within the model, and then I derive a simple wage accounting exercise intended to measure the importance of demand, relative to supply, in explaining longer-run changes in experience differentials within schooling-by-gender groups. Second, I use these tools to describe the joint roles of observed supply shifts and estimated demand shifts in determining changes in the returns to potential experience since the early 1960s. Finally, I discuss the hypothesis that the results presented among women are due exclusively to mis-measurement of actual labor market experience.

Figure 4: Cumulative Changes (times 100) in the Relative Demand for Potential Experience,
Within each Schooling-by-Gender Group



Estimating Relative Demand Shifts and a Simple Wage Accounting Procedure If my empirical measures (including $\widehat{\sigma_e}$) equal their theoretical counterparts, I can re-write equation (4) as:

$$(22) \quad \ln \widehat{w_{sget}} + \frac{1}{\widehat{\sigma_e}} \ln \widehat{g_{set}} = \alpha_{sge} + \alpha_{sgt} + \zeta_{sget} + \varepsilon_{sget}$$

where $\alpha_{sge} = [(\widehat{\sigma_e} - 1) / \widehat{\sigma_e}] \cdot \ln \pi_{sge}$, α_{sgt} captures wage heterogeneity that varies at the level of schooling-by-gender-by-time, $\zeta_{sget} = [(\widehat{\sigma_e} - 1) / \widehat{\sigma_e}] \cdot \ln A_{sget}$, and where the empirical error term ε_{sget} , not explicitly modeled, is included to represent sampling variation and any other sources of variation not captured in the model. I use equation (22) to recover estimates of each demand parameter A_{sget} , which I denote as $\widehat{A_{sget}}$. Then, for each s, g, e and t , I use the estimates $\widehat{\sigma_e}$ and $\widehat{A_{sget}}$ to construct my preferred measure of relative demand within schooling-by-gender aggregates, $\widehat{D_{sget}}$, defined in equation (10).

In Figure 4 and Panels A and B of Table 1, I report changes over time in this measure for each potential experience group within each schooling-by-gender aggregate, and in Panel C of Table 1, I report changes over time in the measure $\widehat{D_{get}} = \widehat{D_{Cget}} - \widehat{D_{Hget}}$ for each gender-by-potential experience group. Before turning to Figure 4 and Table 1 to examine these measures, note that (again, assuming empirical measures equal their theoretical counterparts), differencing equation (4) across potential experience groups, across schooling groups and across time yields the expressions:

$$(23) \quad \Delta \widehat{rel.wage_{sge}} = \frac{1}{\widehat{\sigma_e}} \left[\widehat{\Delta D_{sge}} - \Delta \widehat{rel.supply_{sge}} \right]$$

$$(24) \quad \Delta \widehat{rel.wage_{ge}} = \frac{1}{\widehat{\sigma_e}} \left[\widehat{\Delta D_{ge}} - \Delta \widehat{rel.supply_{ge}} \right]$$

where, with some abuse of notation, I have used the symbol Δ to denote change over time between any two periods.⁴⁹ Equations (23) and (24) imply that the changes in relative wages presented in Table 1 can be decomposed into a supply and a demand component. In the next section, I discuss whether demand shifts played a "large" role, within each gender, in changing experience differentials over the 21-year periods outlined in Table 1; one measure of this role is a simple wage accounting

⁴⁹For example, $\Delta \widehat{rel.wage_{sge}} = \widehat{rel.wage_{sget}} - \widehat{rel.wage_{sget'}}$ for some t and t' .

procedure in which I define the terms:

$$(25) \quad \widehat{demand.share}_{sge} = \frac{|\widehat{\Delta D}_{sge}|}{|\widehat{\Delta D}_{sge}| + |\widehat{\Delta rel.supply}_{sge}|}$$

$$(26) \quad \widehat{demand.share}_{ge} = \frac{|\widehat{\Delta D}_{ge}|}{|\widehat{\Delta D}_{ge}| + |\widehat{\Delta rel.supply}_{ge}|}$$

That is, $\widehat{demand.share}_{sge}$ represents, for workers of schooling group s , gender g and potential experience group e , the change in the term $\widehat{\Delta rel.wage}_{sge}$ that would have occurred over the indicated period if supply had been held fixed, divided by the (gross) change that would have occurred if both supply and demand had shifted (the term $\widehat{demand.share}_{ge}$ is defined similarly). In the table, I present means of these statistics averaged across potential experience groups.

Supply, Demand and Potential Experience Differentials, 1964-2006 Inspection of Figure 4 makes clear that, according to the fitted model, changes in the returns to potential experience among college-educated men appear to be driven by patterns of supply and demand that are qualitatively-different from the patterns driving experience differentials among both female workers and high school-educated men. I discuss these cases in turn.

College Men: Steady Growth in the Relative Demand for Potential Experience Figures plotted in Panel A of Figure 1 and displayed in Panel A of Table 1 are consistent with a large literature that stresses the importance of demand shifts favoring more-experienced workers among males since the early 1960s (Juhn et al., 1993; Murphy and Welch, 1992) and, more-specifically, the original two-part story of Katz and Murphy (1992), who examine experience differentials across the 1963-1987 period. Katz and Murphy hypothesize that smooth growth in the demand for more-experienced workers (potentially driven by skill-biased technological change), combined with fluctuations in the relative supply of the more-experienced (potentially driven by demographic changes such as the entry of the baby boom cohort) combine to explain observed movements in experience differentials.

Consistent with this story, Panel A of Figure 1 shows near-monotonic growth in estimated relative demand for workers with at least 20 years of potential experience, and near-monotonic declines in

the estimated relative demand for workers with fewer than 10 years of potential experience over the entire 1964-2006 period. Overall, relative demand for workers with 20+ years of potential experience increased on average roughly 69% more than the leave-out average, and relative demand for workers with fewer than 10 years of potential experience increased roughly 120% less than the leave-out average over this 42-year period. Inspection of Panel A of Table 1 makes clear that, in the face of this steadily-increasing demand for more-experienced college men, the modest decrease (1964-1985), then large increase (1985-2006) in the relative supply of more-experienced workers created a large increase (1964-1985) then smaller decrease (1985-2006) in the relative wages of more-experienced college men. The wage accounting procedure outlined equations (23) and (25) suggests that, because relative supply movements were relatively-small during the first period, and larger during the second, the role of demand was larger from 1964-1985, accounting for an average of 69% of (gross) movement in relative wages during this period, and smaller from 1985 to 2006, accounting for only 36% of wage movement as increases in relative supply overcame demand increases to push down the wages of the more-experienced.

Women and H.S. Men: Falling, then Rising Relative Demand for Potential Experience

In contrast with college men, and consistent with a large literature on the narrowing of the gender wage gap in the 1980s (Blau and Kahn, 1994; 1997; O'Neill and Polachek, 1993), Figure 4 suggests that the relative demand for more-experienced workers among both high school-educated men and both college- and high school-educated women fell during the 1960s and 1970s and then, beginning in the late 1970s, began to increase. For example, consider the relative demand within each of these schooling-by-gender groups, for workers with fewer than 10 years of potential experience: relative demand series displayed in Panels B, C, and D of Figure 4 first rise in the 1960s and early 1970s; then, in 1979, relative demand within each panel hits an inflection point, and decreases, on average, from 1979 to 2006. Relative demand trends for other potential experience groups also closely-resemble each other across Panels B, C and D of Figure 4: each panel features trend breaks in relative demand that favor workers with at least 20 years of potential experience.

Importantly, trend breaks in relative demand are largest within college-educated women: Panel A of Table 1 shows that the relative demand for college-educated women with 20+ years of experience increased, on average, roughly 4.1% faster annually (or 86% faster total) than the leave-out average among college-educated women across the 1985-2006 period. Finally, note that across this period,

shifts in demand on average accounted for a large share of (gross) changes in experience differentials among relevant groups: 55% among college-educated women, 63% among high school-educated women, and 40% among high school-educated men.

Can Mis-Measurement of Experience Among Women Explain the Results? The results presented above suggest large and economically-significant trend breaks in relative demand favoring more-experienced workers among women beginning in the late 1970s. In this section, I briefly discuss an alternative explanation for these results. According to equation (13), the estimated measures of relative demand for potential experience groups displayed in Figure 4 and Table 1 are interpretable as residuals: changes in these measures represent the increase in relative wages that is unexplained by shifts in relative supply. In the context of the supply and demand model of equations (52), (2) and (3) and wage equation (4), these unexplained wage movements are attributed to (unobserved) shifts in demand across demographic groups that are assumed to be associated with a fixed set of skills. An alternative explanation, however, is that these movements are attributable to changes in composition (changing skills), rather than changing returns, across demographic groups. Although changes in cohort quality are notoriously difficult to separate from shifts in demand across time, here I briefly discuss one particular source of change in cohort quality. One concern with dividing female workers into potential experience groups is that the actual mean labor market experience of women within a given group may not be held fixed over time (O'Neill, 2003a; 2003b). For example, O'Neill and Polachek (1993, Table 4) note that, from 1977 to 1987, while the proportion of (actual) years worked since leaving formal schooling remained roughly constant for men (roughly 92%), this proportion rose for women (roughly from 66% to 71%), and especially among women with 1-15 years of potential experience (roughly from 75% to 83%). Moreover, O'Neill (2003a; 2003b) suggests that these relative increases in the "quality" of younger, less-experienced cohorts of women appear to be important across the entire 1979-2001 period. Can changes in cohort quality driven by shifts in actual labor market experience within given potential experience bins among women explain the patterns observed within women estimated above?

Three facts suggest that this is not likely to be the case. First, as noted above, beginning in 1979, the returns to potential experience appear to increase, not decrease: as such, increases in the relative quality of younger cohorts would suggest that the shifts in relative demand estimated from 1979 onwards among women may under-estimate true relative increase in demand for more-experienced

women. Second, application of the model to high school-educated men - presumably less-affected than women by mis-measurement of labor market experience - reveals demand trends similar to those of women. Third, while panels C and D of Figure 4 display evidence of sharp trend breaks in relative demand - first falling, then rising returns to potential experience among women - increases in female labor force participation among relevant cohorts of women are smooth and monotonic (Altonji and Blank, 1999; Blau and Kahn, 2004; Coleman and Pencavel, 1993; Goldin, 1990; Smith and Ward, 1989).

1.E Changes in the Relative Demand for Female Labor

In this section, I examine the shifts in relative demand for female labor within both college- and high school-educated workers that, together with observed shifts in relative supply, may affect the measurement of the relative demand for college-educated labor (investigated in Section VI). I use the key wage equation of the model, equation (4), to guide my analysis. I first show that, before conducting a supply and demand analysis of wage differentials across schooling-by-gender groups, standard fixed-weight measures of wages and supplies must first be adjusted to reflect the shifts in supply and demand within these groups measured in Section IV. I measure and interpret these adjustments, which I find to be small, in general, relative to movements in standard measures of relative wages and supplies between genders within each schooling group. After this preliminary step, I then follow the approach of Section IV by slowly imposing structure on comparisons of relative wages and supplies in order to examine the potential role of changes in the relative demand for female labor over the 1963-2008 period.

I begin by examining, within both college- and high school-educated workers, the simple comovement of relative wages (the female-male log wage gap) and supplies (the log supply of female, relative to male, labor) over three 15-year periods from 1963-2008. These simple reduced-form comparisons suggest a significant role for shifts in the relative demand for female labor from the late 1970s to the early 1990s. Rather than, as in Section IV, attempting to estimate a particular value of the relevant elasticity of substitution, in the next step I discipline these comparisons by applying the structure of the model, but I investigate a wide range of plausible values for the elasticity of substitution between male and female labor guided by existing estimates in the literature. Finally, I discuss how the results in this section relate to (i) existing findings on shifts in the relative demand for female labor, and (ii) the potential influence of unobserved changes in cohort quality over time.

I.E.1 Adjusting Standard Fixed-Weight Measures of Wages and Supplies

In this section I show that, while in the standard (Katz and Murphy) special case of the model, commonly-used fixed weight measures of wages and supplies are appropriate for use in a supply and demand analysis of gender wage differentials, this is no longer true in the general case of the model, in which shifts in supply and demand *within* schooling-by-gender groups may affect the analysis of shifts in supply and demand *between* these groups. I first derive and interpret adjusted measures of labor supply and log wages of each schooling-by-gender group, and I then measure the differences between these adjusted measures and standard fixed-weight measures.

Deriving and Interpreting Adjusted Measures of Supplies and Wages

Supplies Recall from the previous section that my preferred estimate of σ_e , $\widehat{\sigma}_e = 5.025$, together with estimating equation (22), allows me to recover estimates of each demand parameter A_{sge} , which I denote as \widehat{A}_{sge} . Equation (22) also allows me to estimate each time-invariant relative efficiency weight π_{sge} , which I denote by $\widehat{\pi}_{sge}$.⁵⁰ The empirical measures $\widehat{\sigma}_e$, \widehat{A}_{sge} , and $\widehat{\pi}_{sge}$, together with the supply measure \widehat{g}_{set} can be combined, following the procedure outlined in Section III, to form the empirical analogue of g_{st} from equation (3), which I denote by \widehat{g}_{st} . \widehat{g}_{st} is an empirical index of the labor supplied by workers of schooling group s and gender g at time t .

Inspection of equation (3) shows that, in the standard special case of the model, \widehat{g}_{st} is a simple sum of the quantity of labor supplied by each underlying demographic group (\widehat{g}_{set}), weighted by the effective units of labor supplied in each period by that group ($\widehat{\pi}_{sge}$). In the general case of the model, two changes are made to \widehat{g}_{st} : first, because demand is allowed to shift within each schooling-by-gender group, \widehat{g}_{set} is further weighted by within-group shifts in demand (\widehat{A}_{sge}), so that \widehat{g}_{st} reflects changes in the returns to potential experience in the labor market. Second, because potential experience groups are associated with an elasticity of substitution of $\widehat{\sigma}_e = 5.025$, \widehat{g}_{st} takes the form of a CES aggregate, rather than a simple weighted sum as in the standard case.

⁵⁰In practice, I modify my estimates of π_{sge} from equation (22) in two ways. First, as discussed in detail in Section II, while π_{sge} is intended to capture the *relative* quantity of effective units of labor supplied in each period by a given demographic group, taken literally it represents the product of a common and a relative component. I include a constant in estimation of equation (22), which captures the common component π and allows the term α_{sge} to capture the relative component. Second, after computing each estimate, I follow Autor et al. (2008; Section 6 of Data Appendix) in dividing each value of $\widehat{\pi}_{sge}$ by the value for high school-educated male workers with 10-19 years of potential labor market experience so that values of $\widehat{\pi}_{sge}$ are interpretable relative to this group.

Wages In order to derive an expression for the log wage of these workers, note that the key wage equation, equation (4), can be transformed using the weighting scheme outlined in equation (5) to yield:

$$(27) \quad \widehat{\ln w_{sgt}} - \frac{1}{\widehat{\eta_{sg}}} \left(\frac{\widehat{\sigma_e} - 1}{\widehat{\sigma_e}} \omega_{sgt}^{De} - \frac{1}{\widehat{\sigma_e}} \omega_{sgt}^{Se} \right) = \alpha_{sg} + \alpha_{st} + \frac{\sigma_g - 1}{\sigma_g} \ln A_{sgt} - \frac{1}{\sigma_g} \ln \widehat{g_{st}} + \varepsilon_{sgt}$$

I refer to the LHS of this expression as the "adjusted" measure of the log wage of workers in schooling group s and of gender g at time t ; equation (27) makes clear that this adjusted wage measure is related to the supply of ($\widehat{g_{st}}$) and demand for (A_{sgt}) such workers. Recall that $\widehat{\ln w_{sgt}}$ represents the average composition-adjusted wage of workers in schooling group s and of gender g at time t , weighted using the fixed employment shares $\widehat{\lambda_{sge}}$ defined in Section III and Appendix A.2. Before proceeding with a supply and demand analysis, the LHS of equation (27) makes clear that this standard fixed-weight measure must be adjusted for two linearly-separable effects: the terms ω_{sgt}^{De} and ω_{sgt}^{Se} are the employment-weighted sums of the effects of shifts in relative demand (D) and supply (S), respectively, between potential experience groups (e) within schooling group s and gender g at time t :

$$(28) \quad \omega_{sgt}^{De} = \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \ln \widehat{A_{sgte}} \right)$$

$$(29) \quad \omega_{sgt}^{Se} = \sum_{e=1}^E \left[\widehat{\lambda_{sge}} \cdot \ln \left(\frac{\widehat{g_{set}}}{\widehat{g_{st}}} \right) \right]$$

The term α_{sg} captures the time-invariant (relative) effective units of labor supplied in each period by workers within schooling group s and gender g :

$$\alpha_{sg} = \frac{\sigma_e - 1}{\sigma_e} \cdot \frac{1}{\widehat{\eta_{sg}}} \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \ln \widehat{\pi_{sge}} \right)$$

while the effect α_{st} captures, in addition to wage effects common to all groups captured in the term $(1/\sigma_s) \cdot Y_t$, the effects of supply and demand common to all workers of schooling group s at time t :

$$\alpha_{st} = \frac{1}{\sigma_s} \ln Y_t + \frac{\sigma_s - 1}{\sigma_s} \ln A_{st} - \left[\frac{1}{\sigma_s} - \frac{1}{\sigma_g} \right] \ln s_t$$

and finally ε_{sgt} is the employment-weighted sum of non-modeled variation in log wages:

$$\varepsilon_{sgt} = \frac{1}{\widehat{\eta_{sg}}} \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \varepsilon_{sget} \right)$$

Inspection of the LHS of equation (27) makes clear that, in the standard special case of the model, the adjusted log wage of workers in schooling group s and of gender g at time t simply equals the standard fixed-weight measure $\widehat{\ln w_{sgt}}$, since (i) demand effects, captured in the term ω_{sgt}^{De} , equal 0 by assumption, since $A_{sget} = 1$ for all s, g, e and t , and (ii) shifts in relative supply, captured in the term ω_{sgt}^{Se} , have no effect on log wages since workers are perfect substitutes ($\sigma_e \rightarrow \infty$). Evidence presented in Section IV, however, suggested significant changes over time in relative demand within each schooling group, and an elasticity of substitution of $\widehat{\sigma_e} = 5.025$. In the next section, I measure the extent to which, for each schooling group s and gender g , adjusted measures of log wages (and log supplies) differ from their standard fixed-weight counterparts.

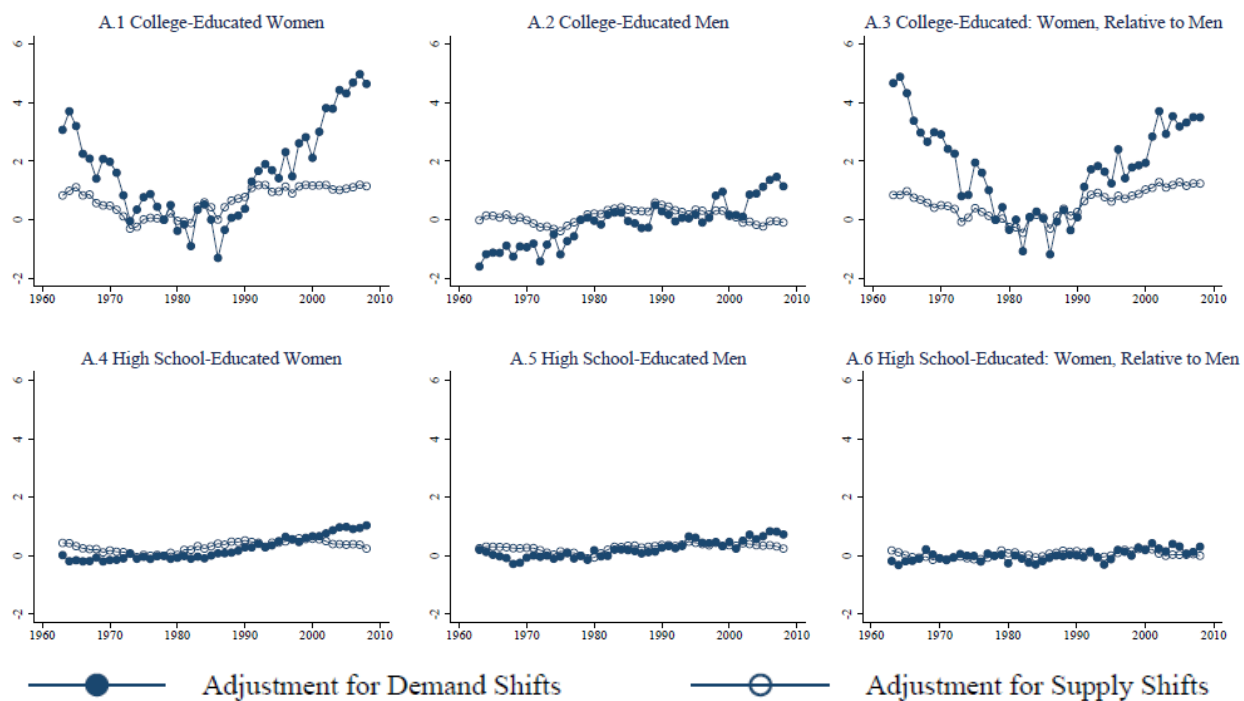
TABLE 4
Measures of Employment Shares (λ_{sge}) and Efficiency Weights (π_{sge})

A. Employment Shares Within Each Schooling Group: Average Across the 1963-2008 Earnings Years (λ_{sge})				
Years of Potential Experience	A.1 College Labor		A.2 High School Labor	
	women	men	women	men
0-9	0.14	0.20	0.12	0.16
10-19	0.10	0.20	0.10	0.17
20-29	0.08	0.15	0.11	0.15
30-39	0.04	0.09	0.08	0.11
Total	0.35	0.65	0.41	0.59

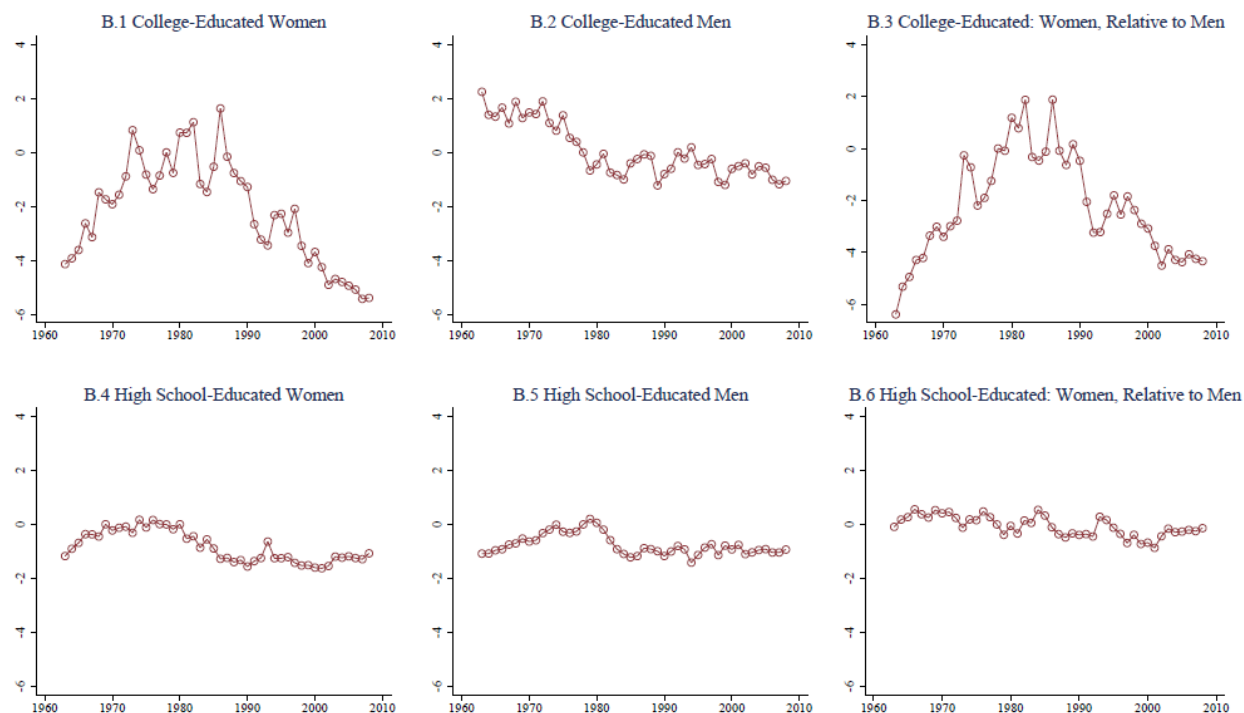
B. Effective Units of Labor Supplied in Each Period, Relative to High School Males with 10-19 Years of Potential Experience (π_{sge})				
Years of Potential Experience	B.1 College		B.2 High School	
	women	men	women	men
0-9	1.15	1.34	0.46	0.73
10-19	1.20	1.87	0.51	1.00
20-29	1.16	1.88	0.53	1.07
30-39	1.03	1.59	0.52	1.04

Figure 5: Measures of Log Wages and Supplies of Schooling-by-Gender Groups:
Adjustments for Shifts in Supply and Demand Between Potential Experience Groups

A. Log Wage Measures: Total Adjustment (multiplied by 100), Decomposed into Supply and Demand Effects



B. Log Supply Measures: Total Adjustment (multiplied by 100)



Measuring Adjustments to Standard Fixed-Weight Measures Figure 5 displays, for each schooling-by-gender group, the difference between the adjusted and standard fixed-weight measures of both log wages and log supplies described above. I begin by describing this difference for measures of log wages.

Log Wages Equation (27) shows that the gap between adjusted and standard measures of log wages can be decomposed into a demand effect of $-(\widehat{\sigma}_e - 1) / (\widehat{\eta}_{sg} \cdot \widehat{\sigma}_e) \cdot \omega_{sgt}^{De}$ and a supply effect of $[1 / (\widehat{\eta}_{sg} \cdot \widehat{\sigma}_e)] \cdot \omega_{sgt}^{Se}$. Two initial patterns stand out in Panels A.1, A.2, A.4 and A.5 of Figure 5. First, while demand effects are important in some panels, shifts in supply within schooling-by-gender groups do not cause more than roughly a 1% difference between adjusted and standard measures of log wages in any panel or period. This results from the relatively high estimated elasticity of substitution between potential experience groups, $\widehat{\sigma}_e = 5.025$, which ensures that movements in relative supplies (demands) have a relatively small (large) impact on wage adjustments. Second, while demand shifts affect the log wages of college-educated workers, they do not significantly affect the wages of high school-educated workers. The reason for this discrepancy lies in equation (28) and in Table 4, which displays the measured employment shares $\widehat{\lambda}_{sge}$ for each of the 16 schooling-by-gender-by-potential experience groups. Inspection of Table 4 reveals that, averaged across the 1963-2008 period, workers with at least 20 years of potential experience represent roughly half of high school workers (47% and 45% among women and men, respectively) while they represent a much smaller minority among the college-educated (35% and 38% among women and men, respectively). Equation (28) clarifies that shifts in relative demand are employment-weighted: since workers with 20+ years of experience represent many fewer than half of college-educated workers, shifts in demand favoring these workers will pull down the standard fixed-weighted wage measure (or, equivalently, will raise the adjusted wage measure) among college-educated workers, but since they represent roughly half of all workers among the high school-educated, such shifts will have little effect on the wage of the high school-educated.

Panels A.1 and A.2 show that adjustments to the log wages of college-educated women and men, respectively, closely-follow shifts in labor demand favoring workers with 20+ years of experience investigated in Section IV of the paper. Among college-educated women, log wages are pushed down by roughly 4%, then up by roughly 5% during the 1963-1978 and 1978-2006 periods, respectively, as labor demand shifted first away from, then towards more-experienced workers. Among college-educated men, log wages are steadily pushed up by roughly 3% as labor demand shifts towards more-experienced

workers.

Log Supplies Unlike log wages, the gap between standard and adjusted log supply indices cannot be decomposed into separate supply and demand components. However, Panels B.1, B.2, B.4 and B.5 of Figure 5 reveal that the measurement gap in log supply is also driven by shifts in demand within schooling-by-gender groups. Inspection of the theoretical analogue of \widehat{g}_{st} , equation (3), demonstrates that, for each potential experience group, the demand parameter \widehat{A}_{sget} modifies the efficiency-weighted quantity of labor supplied, $\widehat{\pi}_{sge} \cdot \widehat{g}_{set}$. Since, as suggested by the discussion above, these quantities are smaller among more-experienced workers, a shift in labor demand favoring more-experienced workers within a given schooling-by-gender group will tend to decrease \widehat{g}_{st} . As a result, adjustments to measures of supply have the opposite pattern observed for log wages: first, among the high school educated, shifts in demand and supply have little effect. Among college-educated women, log supplies are pushed up by roughly 4%, then down by roughly 5%, during the 1963-1978 and 1978-2006 periods as labor demand shifted first away from, then towards more-experienced workers. Among college-educated men, log supplies are steadily pushed down by roughly 2% as labor demand shifts towards more-experienced workers.

I.E.2 Comovement of Relative Wages and Supplies over 15-Year Periods

In this section, I observe simple comovement of female-male log wage and supply gaps within college- and high school-educated workers over three 15-year periods from 1963 to 2008 to examine whether, over each period and within each schooling group, these (relative) quantities and prices moved in opposite directions or whether they moved together, potentially suggesting a role for shifts in relative demand. In contrast to the analysis of Section IV, this analysis is complicated slightly by the adjustments to measures of wages and supplies examined in the previous section; as a result, I first derive the key wage equation that relates the female-male log wage and supply gaps. This allows me to first compare, as in Section IV, simple fixed-weight measures of (relative) prices and quantities, and then to compare the adjusted prices and quantities suggested by the model.

The Key Equation Relating the Female-Male Log Wage and Supply Gaps Differencing equation (27) across gender yields:

$$(30) \quad \widehat{\ln w_{sFt}} - \widehat{\ln w_{sMt}} + \xi_{st}^g = \alpha_s^g + \frac{\sigma_g - 1}{\sigma_g} \ln \left(\frac{A_{sFt}}{A_{sMt}} \right) - \frac{1}{\sigma_g} \ln \left(\frac{\widehat{F}_{st}}{\widehat{M}_{st}} \right) + \varepsilon_{st}^g$$

As in equation (27), I refer to the LHS of equation (30) as the "adjusted" female-male log wage gap within schooling group s at time t . According to the model, it is this adjusted gap, and not the standard gap in fixed-weight log measures, $\widehat{\ln w_{sFt}} - \widehat{\ln w_{sMt}}$, that is simply-related to the relative demand for (RHS term 2) and adjusted supply of (RHS term 3) female labor within schooling group s at time t . The wage adjustment ξ_{st}^g can be linearly decomposed as $\xi_{st}^{gDe} + \xi_{st}^{gSe}$; the adjusted female-male log wage gap, then, is decomposed into three parts: first, the standard fixed-weight log wage gap, then components due to shifts in demand (D) and supply (S), respectively, across potential experience groups (e):

$$(31) \quad \xi_{st}^{gDe} = -\frac{\widehat{\sigma_e} - 1}{\widehat{\sigma_e}} \left(\frac{\omega_{sFt}^{De}}{1 - \widehat{\eta}_s} - \frac{\omega_{sMt}^{De}}{\widehat{\eta}_s} \right)$$

$$(32) \quad \xi_{st}^{gSe} = \frac{1}{\widehat{\sigma_e}} \left(\frac{\omega_{sFt}^{Se}}{1 - \widehat{\eta}_s} - \frac{\omega_{sMt}^{Se}}{\widehat{\eta}_s} \right)$$

Where each term in equations (31) and (32) is defined above.⁵¹ The adjusted female-male log supply gap (RHS term 3 in equation 30) can be simply decomposed into two components:

$$(33) \quad \ln \left(\frac{\widehat{F}_{st}}{\widehat{M}_{st}} \right) = \ln \left(\frac{\widehat{F}_{st}^{\text{F.W.}}}{\widehat{M}_{st}^{\text{F.W.}}} \right) + \left[\ln \left(\frac{\widehat{F}_{st}}{\widehat{M}_{st}} \right) - \ln \left(\frac{\widehat{F}_{st}^{\text{F.W.}}}{\widehat{M}_{st}^{\text{F.W.}}} \right) \right]$$

where the first RHS term of equation (33) represents the standard fixed-weight (F.W.) measure of the female-male gap in log supplies within which, for each $g \in \{F, M\}$, I define:

$$\widehat{g}_{st}^{\text{F.W.}} = \sum_{e=1}^E (\widehat{\pi}_{sge} \cdot \widehat{g}_{set})$$

and where the term in the square brackets of equation (33) represents the difference between

⁵¹Recall from Section III, that $\widehat{\eta}_s = \widehat{\eta}_{sM} = 1 - \widehat{\eta}_{sF}$ represents the average (1963-2008) employment share of men among all workers in schooling group s .

the adjusted and standard measures due to supply and demand adjustments. Finally, to complete description of equation (30), note that the term $\alpha_s^g = \alpha_{sF} - \alpha_{sM}$ reflects time-invariant differences in worker efficiencies, while ε_{st}^g reflects non-modeled wage variation.⁵²

⁵²For brevity, I do not provide explicit expressions for these terms, which are easily-recoverable from equation (27) above.

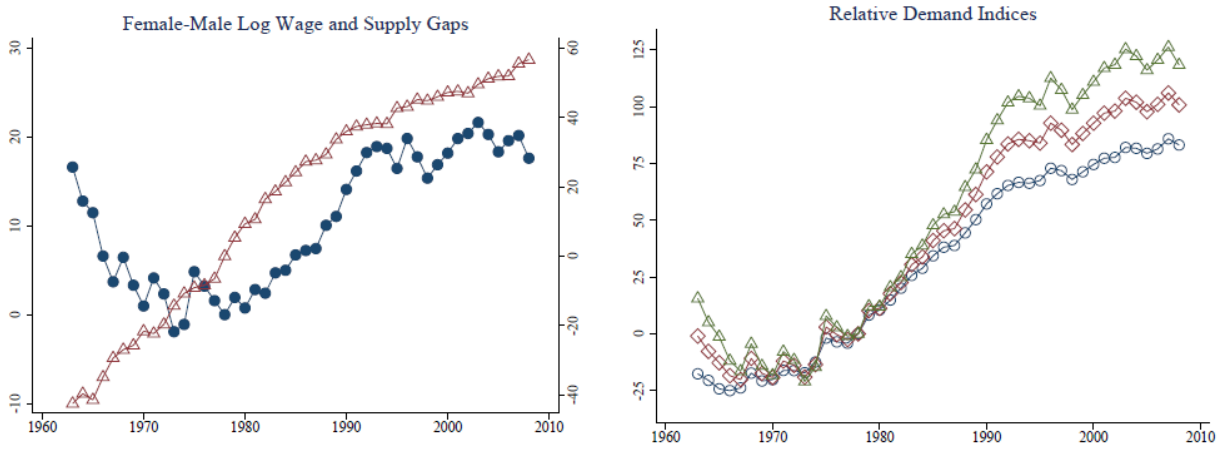
TABLE 5

Annual Changes (times 100) in the Log Relative Wage of, Log Relative Supply of,
and Log Relative Demand for Female Labor Within College and High School Workers, 1963-2008

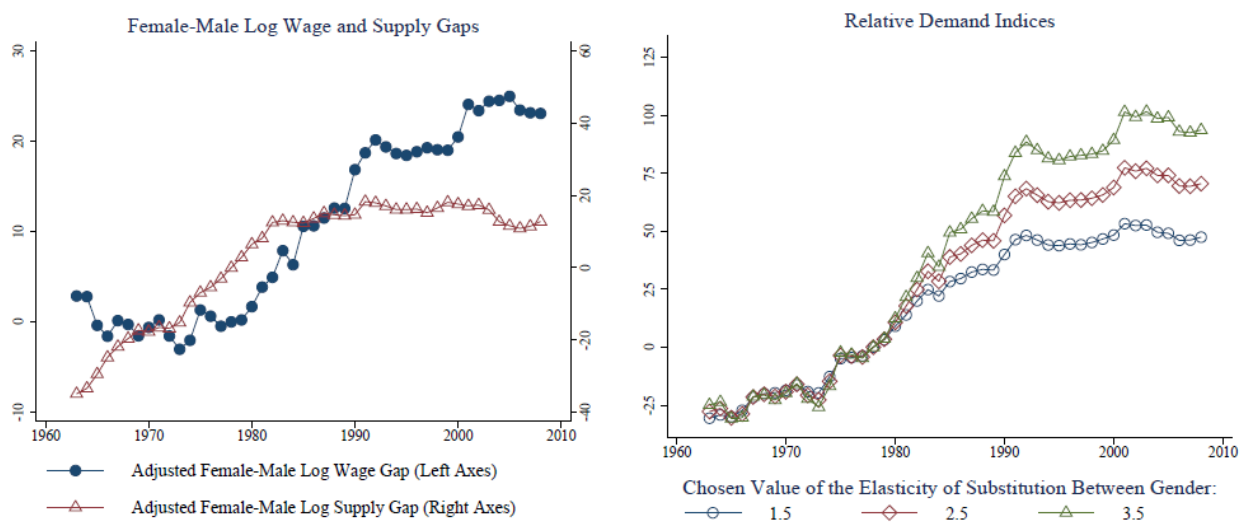
Quantity	A. College Labor			
	1963-1978	1978-1993	1993-2008	1963-2008
Female-Male Adjusted Log Wage Gap				
Standard Fixed-Weight Log Wage Gap	-0.74	1.08	-0.22	0.04
Adjustment: Demand Shifts	-0.31	0.12	0.11	-0.03
Adjustment: Supply Shifts	-0.06	0.06	0.02	0.01
<u>Sum</u> : Total Adjusted Gap	-1.11	1.26	-0.09	0.02
Female-Male Adjusted Log Supply Gap				
Standard Fixed-Weight Log Supply Gap	2.41	2.77	1.30	2.16
Adjustment: Supply and Demand Shifts	0.43	-0.21	-0.07	0.05
<u>Sum</u> : Total Adjusted Gap	2.84	2.55	1.23	2.20
Estimated Relative Demand Index				
$\sigma_g = 1.5$	1.17	4.44	1.10	2.24
$\sigma_g = 2.5$	0.07	5.71	1.01	2.26
$\sigma_g = 3.5$	-1.04	6.97	0.92	2.28
$\sigma_g = 4.5$	-2.15	8.23	0.83	2.30
Quantity	B. High School Labor			
	1963-1978	1978-1993	1993-2008	1963-2008
Female-Male Adjusted Log Wage Gap				
Standard Fixed-Weight Log Wage Gap	-0.19	1.30	0.22	0.44
Adjustment: Demand Shifts	0.01	0.00	0.02	0.01
Adjustment: Supply Shifts	-0.01	0.00	0.00	0.00
<u>Sum</u> : Total Adjusted Gap	-0.19	1.29	0.25	0.45
Female-Male Adjusted Log Supply Gap				
Standard Fixed-Weight Log Supply Gap	2.32	1.12	-0.26	1.06
Adjustment: Supply and Demand Shifts	0.01	0.02	-0.03	0.00
<u>Sum</u> : Total Adjusted Gap	2.33	1.14	-0.29	1.06
Estimated Relative Demand Index				
$\sigma_g = 1.5$	2.04	3.08	0.08	1.73
$\sigma_g = 2.5$	1.85	4.37	0.33	2.18
$\sigma_g = 3.5$	1.66	5.66	0.58	2.63
$\sigma_g = 4.5$	1.47	6.95	0.82	3.08

Figure 6: Cumulative Change (times 100, relative to 1978) in the Log Relative Wage of, Log Relative Supply of, and Relative Demand for Female Labor Within College and High School Workers, 1963-2008

A. College Labor



B. High School Labor



Comovement of Relative Wages and Supplies Table 5 presents changes, over three 15-year periods from 1963-2008, in the adjusted measures of the female-male log wage and supply gaps described in equation (30), as well as decompositions of these changes into the components outlined in equations (31), (32) and (33). To complement this table, the leftmost two panels of Figure 6 plot cumulative changes, relative to the earnings year 1978, in adjusted relative wage and supply measures, while the rightmost panels (A.3, A.6, B.3 and B.6) of Figure 5 plot the adjusted components outlined in equations (31), (32) and (33) for college- and high school-educated labor in each year. I separately-examine, in each 15-year period from 1963-2008, comovement between relative prices and quantities of female labor: as noted above, while negative comovement by itself sheds little light on the role of relative demand, a finding that relative prices and quantities move in the same direction is suggestive of important demand effects if equations (52), (2) and (3) approximate the true structure of aggregate production.

1963-1978 The first period is characterized by quickly-increasing female labor force participation among both the college- and high school-educated: Table 5 documents that, from 1963 to 1978, the adjusted female-male log supply gap increased by roughly 2.8% and 2.3% annually, or 43% and 35% over the entire 15-year period among college and high school labor, respectively. Figure 6 shows that this increase was nearly-linear across time among both groups. The log wage gap, however, behaved quite differently across these groups: while the adjusted log wage gap among the high school-educated was roughly constant from 1963-1978, the adjusted log wage gap among the college-educated decreased significantly, falling close to roughly 16% over this period. Inspection of Table 5 and Panel A.3 of Figure 5 reveals that while roughly 30% of this drop is due to shifts in demand towards more-experienced workers among college-educated women documented in Section IV, the remainder (70% of the total, or an 11% decrease in the log wage gap) is due, in an accounting sense, to a drop in the standard fixed-weight measure of the log gender wage gap among the college-educated.⁵³ Overall, joint movement of relative wages and supplies across this period is not, by itself, indicative of relative demand increases, as relative supplies of female labor increased quickly among both schooling groups,

⁵³This steep decrease is confirmed in Acemoglu and Autor (2011); numbers reported in their Table 1a suggest that the log female-male log wage gap among workers with 16+ years of education decreased by roughly 9.2% from 1963-1979, which is similar to the roughly 11% drop from 1963-1978 I report here. These findings are not consistent, however, with the finding of similar paths in the returns to college among males and females across this period (Katz and Murphy 1992; Figure 1). This discrepancy may be attributable to methodological differences between studies.

potentially playing a role in the large (college-educated) and mild (high school-educated) decreases in the gender wage gaps observed across this period.

1978-1993 Shifts in relative demand favoring female workers, however, appear to be strongly suggested across the next 15-year period. Consistent with a large literature documenting the narrowing of the gender wage gap during the 1980s (Blau and Kahn, 1994; 1997; O'Neill and Polachek, 1993; Smith and Ward, 1989) I find that, between 1978 and 1993, the adjusted female-male log wage gap increased by roughly 1.3% annually, or roughly 20% across the entire period, among both the college- and high school-educated. During the early 1980s, however, patterns of relative supplies begin to starkly differ across schooling groups. While the relative supply of female labor maintained its rapid increase among college-educated workers, increasing by roughly 38% across the 1978-1993 period, Figure 6 demonstrates a clear negative trend break in 1982 among the high school-educated. This trend-break is consistent with evidence on historical educational investment decisions: Goldin and Katz (2008, e.g., Figure 7.1) document that the marked slowdown in the college graduation rate among birth cohorts of the 1950s and 1960s was much more-pronounced among men; Figure 5 shows that the trend in log relative supply of female labor is quickly-increasing across the entire 1963-2008 period among the college-educated, but features a sharp slowdown among the high school-educated. Indeed, fitting the relative supply series among college-educated workers plotted in the top-left panel of Figure 6 with a linear trend is associated with an R-squared of 0.956, while fitting the relative supply series among high school-educated workers with a simple linear spline that features a single trend break in 1982 is associated with an R-squared of 0.983. Overall, comparisons across the 1978-1993 period suggest that simultaneously-increasing relative prices and quantities of female labor may be reconciled by increases in relative demand, especially among the college-educated, for whom the relative supply of female labor strongly-increased over the entire period.

1993-2008 While trends in the relative supply of female labor discussed above continued into the 1993-2008 period, with the adjusted female-male log supply gap increasing by roughly 1.2% and -0.3% annually among the college- and high school-educated, respectively, trends in relative wages change sharply for both groups in the early 1990s. Consistent with a more-recent literature on the gender wage gap (Blau and Kahn, 2004), compression of the gender wage gap slowed considerably among both groups in 1993: rather than continuing to swiftly-increase among both groups, the adjusted

female-male log wage gap decreased by roughly 1.3% and increased by only 3.7% across the entire 15-year period from 1993-2008 among college- and high school-educated workers, respectively.

A Role for Demand Shifts? Overall, inspection of Table 5 and the wage and supply series from Figures 6 and 7 appear to suggest a large role for shifts in relative demand favoring female workers among both schooling groups, and particularly among the college-educated, during the 1978-1993 period. A stark difference in the pattern of supplies across schooling groups is also potentially instructive: while the relative supply of female labor among the college educated saw strong, nearly-linear growth across the entire 1963-2008 period, relative supply among the high school-educated saw a sharp trend decrease that coincided with the overall slowdown in the college equivalent share of the labor force in 1982. To the extent that supply increases push down the relative wage of women (an effect governed, in the model, by the elasticity σ_g), these patterns are potentially-suggestive of a larger role for relative demand increases among the college-educated than among the high school-educated after 1982. In the next section, I investigate these hypotheses formally by applying the structure of the model to comparisons of wages and supplies in order to understand the potential role of demand.

I.E.3 Estimated Changes in the Relative Demand for Female Labor

The reduced-form comparisons of the previous section suggest that the simultaneous narrowing of the gender wage gap and relative expansion of the female labor force within both college- and high school-educated workers between the late 1970s and the early 1990s may potentially be reconciled by increases in the relative demand for female labor within each schooling group. In this section, I discipline these comparisons by applying the structure of production assumed in equations (52), (2) and (3). In contrast to the empirical approach of Section IV, in which I exploited plausibly-exogenous variation in relative labor supply to estimate a preferred value for the relevant elasticity of substitution, in this section I take a different approach. Because variation in the relative supply of female labor is likely to be driven in large part by changes along the labor force *participation* margin and is thus, even in the short run, likely related to changes in the relative demand for female labor, in this section (and in Section VI of the paper) I investigate changes in the relative demand for female labor implied by a wide range of plausible values for the elasticity of substitution σ_g .

This section proceeds in four parts. In the first part, I outline relevant elasticity estimates from existing studies that motivate the range of values for σ_g that I choose to investigate. Second, I use

these values, together with the key wage equation of the model, equation (4), to estimate indices of the relative demand for female labor within both college- and high school-educated workers from 1963-2008. Consistent with the suggestive evidence presented above, I find that for every value of σ_g investigated, the fitted model suggests large trend breaks in relative demand favoring female workers within both schooling groups, and particularly within the college-educated. These results are consistent with, but contain important differences from, estimates of shifts in the relative demand for female labor resulting from changes in employment across industry-by-occupation cells in the literature: in the penultimate section, I discuss these differences. Finally, I discuss whether variation in unobserved worker quality across cohorts can explain the results of this section.

Relevant Estimates from the Literature Potentially due to the difficulty in isolating exogenous variation in the relative supply of female labor, there are surprisingly few existing estimates of elasticities of substitution between male and female labor (Hamermesh, 1996).⁵⁴ To my knowledge, an estimate of the parameter σ_g - the (Allen) elasticity of substitution between male and female equivalent labor within college and high school aggregates - has yet to be estimated.⁵⁵ A small group of papers estimate a closely-related parameter: the (Allen) elasticity of substitution between all male and female labor. I use these estimates to guide the values of σ_g that I will investigate.

Two papers estimate this elasticity by fitting a model directly to national time-series data. Weinberg (2000) estimates an elasticity of 2.4 using CPS data from the U.S. across the period 1970 to 1994, while Lewis (1985) estimates an elasticity of 2.29 using yearly data on the Australian economy from 1975-1981. Layard (1982) takes a different approach to identification: he assumes that manufacturing firms in Britain are price takers, and thus that industry-wide relative quantity movements over time (among manual workers) in this sector trace out an industry-level relative demand curve. Using time series data from 1949-1969, he estimates an elasticity of substitution of 2.0. Finally, Acemoglu and

⁵⁴As discussed in more detail in Section II, there is disagreement in the literature regarding *which types* of men and women should be modeled as direct substitutes in the literature (and, hence, which elasticities of substitution should be estimated). Since I estimate the elasticity of substitution between men and women of the same schooling group, in this section I examine elasticity estimates from the literature that are most-directly related to the associated parameter of interest, σ_g .

⁵⁵Although Katz and Murphy (1992) compute (absolute) industry-by-occupation-based measures of demand shifts for groups very similar to the aggregates F_{st} and M_{st} within aggregates very similar to schooling groups s , and although Blau and Kahn (1997) compute (relative) industry-by-occupation-based measures of shifts in demand for female labor within regression-delineated "predicted" skill groups, neither of these papers estimates an elasticity of substitution between female and male labor.

Autor (2004) use cross-state variation in the relative supply of female labor due to World War II to estimate a (state-level) elasticity of substitution between male and female labor. These authors present a number of different approaches to estimating this elasticity; estimates range from roughly 1.5 to roughly 4.2, with their preferred estimates centered around 3.

In my empirical work, I examine a range of values for σ_g that spans these estimates: I estimate the implied path of relative demand for female labor within both college- and high school-educated workers given values of σ_g between 1.5 and 4.5.⁵⁶ In the next section, I first show how, given a choice for σ_g , I can use the key wage equation of the model to estimate shifts in relative demand. I then examine changes over time in the relative demand for female labor.

Estimated Changes in Relative Demand If my empirical measures equal their theoretical counterparts, I can manipulate equation (4) to yield:

$$(34) \quad \widehat{\ln w_{sgt}} - \frac{1}{\widehat{\eta_{sg}}} \left(\frac{\widehat{\sigma_e} - 1}{\widehat{\sigma_e}} \omega_{sgt}^{De} - \frac{1}{\widehat{\sigma_e}} \omega_{sgt}^{Se} \right) + \frac{1}{\widehat{\sigma_g}} \ln \widehat{g_{st}} = \alpha_{sg} + \alpha_{st} + \zeta_{sgt} + \varepsilon_{sgt}$$

where $\zeta_{sgt} = [(\widehat{\sigma_g} - 1) / \widehat{\sigma_g}] \cdot \ln A_{sgt}$, and where all other terms are defined as in equation (27). I use a given choice of $\widehat{\sigma_g}$, combined with estimates of all other parameters in equation (34) to estimate the parameter A_{sgt} for each schooling group s , gender g and time t . Then, using equation (9), I use this estimate, denoted by $\widehat{A_{sgt}}$, together with the given value of $\widehat{\sigma_g}$, to construct for each schooling group s and each time t the index of within-schooling group relative demand for female labor, $\widehat{D_{st}}$. In Table 5 and in the rightmost panels of Figure 6, I present changes over time in $\widehat{D_{st}}$ for both college- and high school-educated workers using the range of choices for $\widehat{\sigma_g}$ described above. Below, I find evidence of important breaks in the trend growth of $\widehat{D_{st}}$ in the late 1970s; I therefore separate my discussion of the entire 1963-2008 period into the first 15 years (1963-1978) and the next 30 years (1978-2008).

1963-1978 The relative supply of female labor increased quickly among both schooling groups from 1963 to 1978, but the paths of relative wages differed dramatically, as adjusted relative wages decreased roughly 14% more among the college-educated than among the high school-educated. Es-

⁵⁶Blau and Kahn (1997) take a similar approach, examining a range of values for an elasticity of substitution between male and female labor of between 1 and 3.

timated shifts in relative demand reflect this difference: for each value of σ_g , relative demand for female labor is measured to have grown more-slowly among the college-educated across this period. For example, a chosen value of $\sigma_g = 2.5$ (the value closest to most of the existing estimates in the literature discussed above) implies that the index of relative demand for female labor increased by roughly 0.07% and 1.85% annually, or 1% and 28% over the entire 1963-1978 period among college- and high school-educated workers, respectively.

1978-2008 Across the next 30 years, however, this trend difference between college and high school labor reverses. As documented in detail above, inspection of Table 5 and Figure 6 suggest that while trend growth in the relative female wage is relatively similar within both schooling groups across the 1978-2008 period, trend growth in the relative supply of female labor is not: in particular, trend growth among high school-educated workers features a sharp slowdown in 1982. Estimated relative demand indices presented in Table 5 and Figure 6 make clear that, to an extent governed by the parameter σ_g , the continued increase in the relative supply of female labor within the college-educated requires a greater increase in the relative demand for female labor among college-educated workers than among high school-educated workers in order to reconcile the time series of relative wages and supplies.⁵⁷

As suggested by the comparisons presented above, I find large increases in the relative demand for female labor, within both schooling groups, during the 1978-1993 period, followed by positive, but much smaller increases thereafter: for example, the value $\sigma_g = 2.5$ implies indices of relative demand for female labor that increase by roughly 86% and 67% (15% and 5%) over the 1978-1993 (1993-2008) period among college- and high school-educated workers, respectively. Note that, across both of these 15-year periods, and in stark contrast to the 1963-1978 period, for every considered value of σ_g , the rate of increase in the relative demand for female labor is greater within college labor than within high school labor.

⁵⁷Less substitutability between female and male labor, associated with smaller values of σ_g , implies a larger impact of changing relative supply on the relative wage and, in turn, implies that larger shifts in relative demand A_{sFt}/A_{sMt} are required to reconcile the simultaneous increase in the relative wage and supply of female labor within both schooling groups after 1978. However, note that the relative demand index \widehat{D}_{st} , defined in equation (9), takes the form $(\widehat{\sigma}_g - 1) \cdot \ln(\widehat{A_{sFt}}/\widehat{A_{sMt}})$, which adjusts for differences in σ_g . Thus, in Table 5 and Figure 6 it is not the case that measured increases in \widehat{D}_{st} across the 1978-2008 period are decreasing in $\widehat{\sigma}_g$.

Comparisons with Existing Estimates The finding of large trend breaks in relative demand favoring female labor beginning in the late 1970s, followed by slower increases beginning in 1993 is, as discussed in Section II, broadly-consistent with the literature on the evolution of the gender wage gap, which consistently estimates large changes in the returns to female-specific demographic characteristics in the labor market and suggests that shifts in demand favoring women may play a large role in these changing returns (Blau and Kahn, 1994; 1997; 2004; Katz and Murphy, 1992; O’Neill and Polachek, 1993; Smith and Ward, 1989). By contrast, the notion that, over the period 1978-2008, these shifts were larger among the college-educated than among the high school-educated (i) is the subject of far less empirical research, and (ii) is consistent with, but stands in contrast to existing estimates of labor demand shifts derived from industry-by-occupation-level employment changes during the 1980s, which suggests larger increases in the relative demand for female labor within the high school-educated than within the college-educated. Katz and Murphy (1992, Table 6) compute demand shifts derived from employment changes across industry-by-occupation cells across the 1967-1987 (1979-1987) period that are roughly 4.3% and 16.1% (-2.5% and 6.5%) larger for women than men among the college- and high school-educated, respectively. Blau and Kahn (1997; Table 3) follow this work and find similar results, suggesting that, from 1979 to 1988, industry-by-occupation-driven demand shifts resulted in a slight decrease in the relative demand for female labor among the most-skilled workers, but an increase within other skill groups.

Two factors may aid in explaining the discrepancy between these results. First, while demand shifts inferred from comparisons of relative wages and supplies in principle capture all sources of variation in the relative demand for female labor - potentially including, for example, difficult-to-measure but potentially-important demand sources such as changes in the attitudes of employers regarding hiring female workers - demand shifts computed from gross movements in industry-by-occupation cells capture a more-limited set of sources of demand variation. For example, the large decreases in (male-intensive) basic manufacturing work and production jobs that appear as likely proximate causes of relative increases of demand for female labor within high school-educated workers (Katz and Murphy, 1992; Table 5) likely capture only a small portion of total variation in gender-specific shifts in labor demand.⁵⁸ Second, while demand shifts measured in the literature attempt to

⁵⁸For example, in their analysis of increases in the demand for college- relative to high school-equivalent workers, Katz and Murphy note that demand shifts computed from employment changes in industry-by-occupation cells capture about one third of the increase in relative demand inferred by comparing relative wages and supplies.

capture raw shifts in relative labor demand, recall that the demand parameters A_{sgt} , which form the basis of the relative demand index \widehat{D}_{st} , are net of measured demand shifts across potential experience groups within each schooling-by-gender group. For example, a comparison of Panels A.3 and A.6 of Figure 5 show that while the (fixed-weight) female-male log wage gap within the college-educated is adjusted upwards by roughly 5% during the 1978-2008 period due to the effects of supply and demand shifts within schooling-by-gender groups, similar adjustments are very small within high school-educated workers.

Can Unobserved Changes in Cohort Quality Explain the Results? Throughout this paper, I make the common assumption that the unobserved "quality," or unobserved skill, of demographic groups under consideration is approximately fixed; however, if this is not true, it is possible that cross-group changes in the *content* of skill (due to cross-cohort compositional changes over time), rather than the *returns* to skill (due to supply and demand shifts over time) may explain some fraction of the observed changes in the gender differentials investigated across the 1963-2008 earnings years. In a recent paper, Mulligan and Rubenstein (2008) explicitly models the selection bias associated with women's decision to enter the labor force (Roy, 1951; Gronau, 1974; Heckman, 1974), and finds that a large fraction of the narrowing of the gender wage gap is explained by changes in cohort quality. A large fraction of the work discussed above, by contrast, begins by assuming that, after accounting for compositional changes across (potential) labor market experience and schooling, all other dimensions of worker quality are held fixed (e.g., Blau and Kahn, 1994; 1997; O'Neill and Polachek, 1993). Here, I follow these authors: by applying the structure of aggregate production outlined in equations (52), (2) and (3) and, more-specifically, by (i) comparing genders within schooling groups, thus holding fixed (broad) schooling-based compositional changes, (ii) using the wage measures of equation (6) that account for changes in composition across potential experience groups within each schooling-by-gender group, and (iii) adjusting for changes in the returns to potential experience within each schooling-by-gender group using equation (30), I hope to control for the majority of changes in cohort quality across time. If, as is plausible, however, the compositional changes along the *observed* schooling and labor market experience margins that Blau and Kahn (1997) and O'Neill and Polachek (1993) find explain roughly 2/3 and 1/2 of the narrowing of the gender wage gap during the 1980s are associated with simultaneous but *unobserved* relative increases in worker quality among women, measures of shifts in

relative demand may indeed be affected.⁵⁹

1.F Changes in the Relative Demand for the College-Educated

In Sections IV and V above, I examined shifts in supply and demand within college- and high school-educated workers. In this section, I examine how these supply and demand shifts affect changes over time in the key measure of skill-biased technological change in the standard model: the relative demand for college- versus high school-educated labor. As in previous sections, I use equation (4) to guide my analysis: I use this equation to show that, in the general case of the model, standard fixed-weight measures of the relative wage and supply of college-educated labor must first be adjusted for shifts in supply and demand *within* schooling groups before proceeding with an analysis of shifts in supply and demand *between* schooling groups. I then show that, in the general model, the relative demand for college-educated labor equals the standard measure plus a set of adjustments for these supply and demand shifts that can be explicitly derived from the estimates of Sections IV and V of the paper.

As in previous sections, I slowly apply the structure of the model. First, for a broad range of parameter values, I examine changes over time in standard fixed-weight measures of the relative wage and supply of the college-educated, and I examine adjustments to these measures that are suggested by the model: while the adjusted log relative supply series is very similar to the standard fixed-weight benchmark, growth in the adjusted college wage premium is significantly more-rapid than its fixed-weight counterpart, especially since the early 1980s, suggesting that shifts in supply and demand within schooling groups have depressed the composition-adjusted college wage premium over this period. I examine in detail the sources of this wage effect, and consider the hypothesis that shifts in relative demand within schooling groups - favoring women and favoring workers with more potential experience among women - beginning in the late 1970s and examined in detail in Sections IV and V above, are drivers of the effect.

Taken together, these wage and supply effects suggest that growth in the relative demand for college-educated workers may have been more-rapid than standard measures imply, especially since

⁵⁹One standard method of distinguishing between changes over time in cohort composition versus changes in labor market returns is to examine changes over time in prices of interest within cohorts: a classic example is Juhn et al. (1993), who find that changes in the composition of unobserved skills are unlikely to explain a large fraction of increasing residual wage inequality. However, a similar test in the current context - examining within-cohort change in the log female-male wage gap - is not easily-interpretable, given the changes over time in the returns to potential labor market experience measured in Section IV.

the early 1980s; in the main empirical section of the paper, I examine in detail the standard relative demand measure and the adjustments to this measure implied by the generalized model for a wide range of values of the elasticities of substitution between both gender (σ_g) and schooling (σ_s). Finally, I estimate the elasticity σ_s using the relative wage equation implied by the model, and I compare estimates of this parameter implied by imposing the assumptions of the standard special case of the model to those implied by relaxing these assumptions.

I.F.1 College-Educated Labor: Relative Wages, Supplies and Demands

In this section, I manipulate the key wage equation of the model, equation (4), to yield two expressions that will guide my empirical work. First, I derive a wage equation that relates the relative wage of college-educated workers (the college wage premium) to the relative supply of the college-educated: the resulting expression is a generalization of the standard Katz and Murphy (1992) wage equation in which the standard fixed-weight measures of the college wage premium and the relative supply of the college-educated are adjusted by terms that can be explicitly derived using values of model parameters and the supply and demand shifts estimated in Sections IV and V of the paper. Second, I show that, as in the standard Katz and Murphy equation, this wage equation yields a residual measure of growth in the relative demand for college-educated workers that implicitly compares increases in the (adjusted) college wage premium to shifts in the (adjusted) relative supply of college workers. I show that relative demand in the general model can be expressed as the standard measure plus a set of adjustments that can be meaningfully-decomposed into effects due to shifts in supply and demand within schooling groups.

The (Katz Murphy) Relative Wage Equation Assuming empirical measures equal their theoretical counterparts, equation (4) can be weighted using equation (5) to yield an expression for the (composition-adjusted) log wage of schooling group s at time t :

$$(35) \quad \widehat{\ln w_{st}} + \xi_{st} = \alpha_s + \alpha_t + \frac{\sigma_s - 1}{\sigma_s} \ln A_{st} - \frac{1}{\sigma_s} \ln \hat{s}_t + \varepsilon_{st}$$

Here $\widehat{\ln w_{st}}$ is the standard fixed-weight measure of the log wage of schooling group s at time t defined in equation (5), and \hat{s}_t is the measure of the log supply of workers of schooling group s at time

t , defined in Section III and an empirical analogue of the schooling aggregate s_t defined in equation (2) that can be decomposed as:

$$(36) \quad \ln \widehat{s}_t = \ln \widehat{s}_t^{\text{F.W.}} + \left[\ln \widehat{s}_t - \ln \widehat{s}_t^{\text{F.W.}} \right]$$

where $\widehat{s}_t^{\text{F.W.}}$ denotes the standard fixed-weight empirical analogue of equation (2) that is implied under the assumptions of the standard model; that is, where

$$\widehat{s}_t^{\text{F.W.}} = \sum_g \sum_{e=1}^E (\widehat{\pi_{sge}} \cdot \widehat{g_{set}})$$

According to equation (35), after adjusting for the term ξ_{st} , the standard fixed-weight log wage measure $\widehat{\ln w_{st}}$ is simply-related by the elasticity σ_s to the demand for, and (adjusted) supply of, labor of schooling group s at time t . The term ξ_{st} can be expressed as the simple sum of four supply (S) and demand (D) terms: two that capture shifts across gender (g), and two that capture shifts across potential experience groups (e) within schooling groups:

$$(37) \quad \xi_{st}^{Dg} = -\frac{\widehat{\sigma_g} - 1}{\widehat{\sigma_g}} \omega_{st}^{Dg}$$

$$(38) \quad \xi_{st}^{Sg} = \frac{1}{\widehat{\sigma_g}} \omega_{st}^{Sg}$$

$$(39) \quad \xi_{st}^{De} = -\frac{\widehat{\sigma_e} - 1}{\widehat{\sigma_e}} (\omega_{sFt}^{De} + \omega_{sMt}^{De})$$

$$(40) \quad \xi_{st}^{Se} = \frac{1}{\widehat{\sigma_e}} (\omega_{sFt}^{Se} + \omega_{sMt}^{Se})$$

where the terms in equations (39) and (40) are the weighted sums of the relative demand and supply effects defined in equations (28) and (29), respectively, and where

$$(41) \quad \omega_{st}^{Dg} = (1 - \widehat{\eta}_s) \cdot \ln \widehat{A_{sFt}} + \widehat{\eta}_s \cdot \ln \widehat{A_{sMt}}$$

$$(42) \quad \omega_{st}^{Sg} = (1 - \widehat{\eta}_s) \cdot \ln \left(\frac{\widehat{F_{st}}}{\widehat{s_t}} \right) + \widehat{\eta}_s \cdot \ln \left(\frac{\widehat{M_{st}}}{\widehat{s_t}} \right)$$

To summarize, equation (35) suggests that, prior to an investigation of shifts in supply and demand between schooling groups, the standard fixed-weight measures of log wages and supplies, $\widehat{\ln w_{st}}$ and $\ln \widehat{s_t}^{\text{F.W.}}$, respectively, must be adjusted for shifts in supply and demand within group s . Equations (37) through (42) make clear that, for wage measures, these adjustments can be decomposed into four components, each of which is an employment-weighted sum of the supply and demand shifts investigated in Sections IV and IV of the paper, while equation (36) shows that adjusted supply can be decomposed into the standard fixed-weight measure plus the adjustment. Finally, to complete the description of equation (35), the fixed effects and error terms take the form:

$$\alpha_s = \frac{\widehat{\sigma_e} - 1}{\widehat{\sigma_e}} \cdot \sum_g \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \ln \widehat{\pi_{sge}} \right)$$

$$\alpha_t = \frac{1}{\sigma_s} \ln Y_t$$

$$\varepsilon_{st} = \sum_g \sum_{e=1}^E \left(\widehat{\lambda_{sge}} \cdot \varepsilon_{sget} \right)$$

That is, α_s is a simple weighted sum of the effective units of labor supplied in each period by each demographic group of workers within schooling group s , α_t is a term that affects all demographic groups equally at a given time (and will thus drop from all relative wage expressions), and the error term ε_{st} reflects the weighted sum of all non-modeled wage variation within schooling group s at time t .

The generalization of the Katz and Murphy (1992) relative wage equation is obtained by simply differencing equation (35) across schooling groups:

$$(43) \quad \left(\widehat{\ln w_{Ct}} - \widehat{\ln w_{Ht}} \right) + \xi_t = \alpha + \frac{\sigma_s - 1}{\sigma_s} \ln \left(\frac{A_{Ct}}{A_{Ht}} \right) - \frac{1}{\sigma_s} \ln \left(\frac{\widehat{C_t}}{\widehat{H_t}} \right) + \varepsilon_t$$

where, at time t , the fixed-weight college wage premium in the curly brackets on the LHS is adjusted by the term $\xi_t = \xi_t^{Dg} + \xi_t^{Sg} + \xi_t^{De} + \xi_t^{Se}$, where $\xi_t^{Dg} = \xi_{Ct}^{Dg} - \xi_{Ht}^{Dg}$ and the terms ξ_t^{Se} , ξ_t^{Dg} , ξ_t^{Sg} are defined similarly, where the constant $\alpha = \alpha_C - \alpha_H$ reflects the gap between the effective units of labor supplied in each period by college- versus high school-educated workers, where the term $\varepsilon_t = \varepsilon_{Ct} - \varepsilon_{Ht}$ reflects variation in the log wage of the underlying demographic groups that is not captured in the assumed structure of aggregate production in equations (52), (2) and (3) and where adjusted log supply can be decomposed as the standard fixed-weight measure and adjustments due to shifts in supply and demand within schooling groups:

$$(44) \quad \ln \left(\frac{\widehat{C_t}}{\widehat{H_t}} \right) = \ln \left(\frac{\widehat{C_t}^{F.W.}}{\widehat{H_t}^{F.W.}} \right) + \left[\ln \left(\frac{\widehat{C_t}}{\widehat{H_t}} \right) - \ln \left(\frac{\widehat{C_t}^{F.W.}}{\widehat{H_t}^{F.W.}} \right) \right]$$

I return in detail to the measurement of adjustments to standard fixed-weight measures of log wages and supplies in equations (35) and (43) below; first, I derive the measure of relative demand for college-educated labor implied by equation (43).

The Relative Demand for College-Educated Labor Given a value for the elasticity of substitution between schooling groups, denoted $\widehat{\sigma}_s$ and assumed to equal the true value σ_s , equation (43) can be manipulated to yield:

$$(45) \quad \begin{aligned} \widehat{D_t} &= \widehat{\sigma}_s \cdot \left(\widehat{\ln w_{Ct}} - \widehat{\ln w_{Ht}} \right) + \ln \left(\frac{\widehat{C_t}^{F.W.}}{\widehat{H_t}^{F.W.}} \right) \\ &\quad + \widehat{\sigma}_s \cdot \left(\xi_t^{De} + \xi_t^{Se} + \xi_t^{Dg} + \xi_t^{Sg} \right) \\ &\quad + \left[\ln \left(\frac{\widehat{C_t}}{\widehat{H_t}} \right) - \ln \left(\frac{\widehat{C_t}^{F.W.}}{\widehat{H_t}^{F.W.}} \right) \right] \\ &\quad - \widehat{\sigma}_s \cdot (\alpha + \varepsilon_t) \end{aligned}$$

That is, the standard measure of relative demand for college workers, defined in equation (8) and represented in the LHS of equation (45), can be decomposed into several components. The first line of the RHS of equation (45) is the standard measure of the relative demand for college-educated workers, and is the empirical analogue to equation (11), which can be interpreted as that component of the (fixed-weight) relative wage that is unexplained by the value of (fixed-weight) relative supply. The subsequent lines of the RHS of equation (45), however, make clear that in the general case of the model, relative demand *between* schooling groups must be adjusted for the effects of shifts in supply and demand *within* schooling groups on the standard fixed-weight measures of log wages (line 2) and log supplies (line 3).⁶⁰

For intuition, suppose that the standard assumption of common demand shifts - that the relative demand parameters A_{sgt} and A_{sget} defined in equations (2) and (3) equal 1 for every s , g , e and t - is not literally true, as I have argued in Sections IV and V. In this case (and, assuming perfect substitution within schooling groups for simplicity), imposing this assumption mechanically forces lines 2 and 3 of equation (45) to equal 0, meaning that these effects will be captured by the term ε_t , so that the relative demand for college-educated workers \widehat{D}_t will be mis-measured. In the following sections, I investigate the empirical relevance of this potential mis-measurement.

⁶⁰The final line of equation (45) consists of the constant term α , which falls out of any time-differenced expression, and the term ε_t , which captures any non-modeled wage variation.

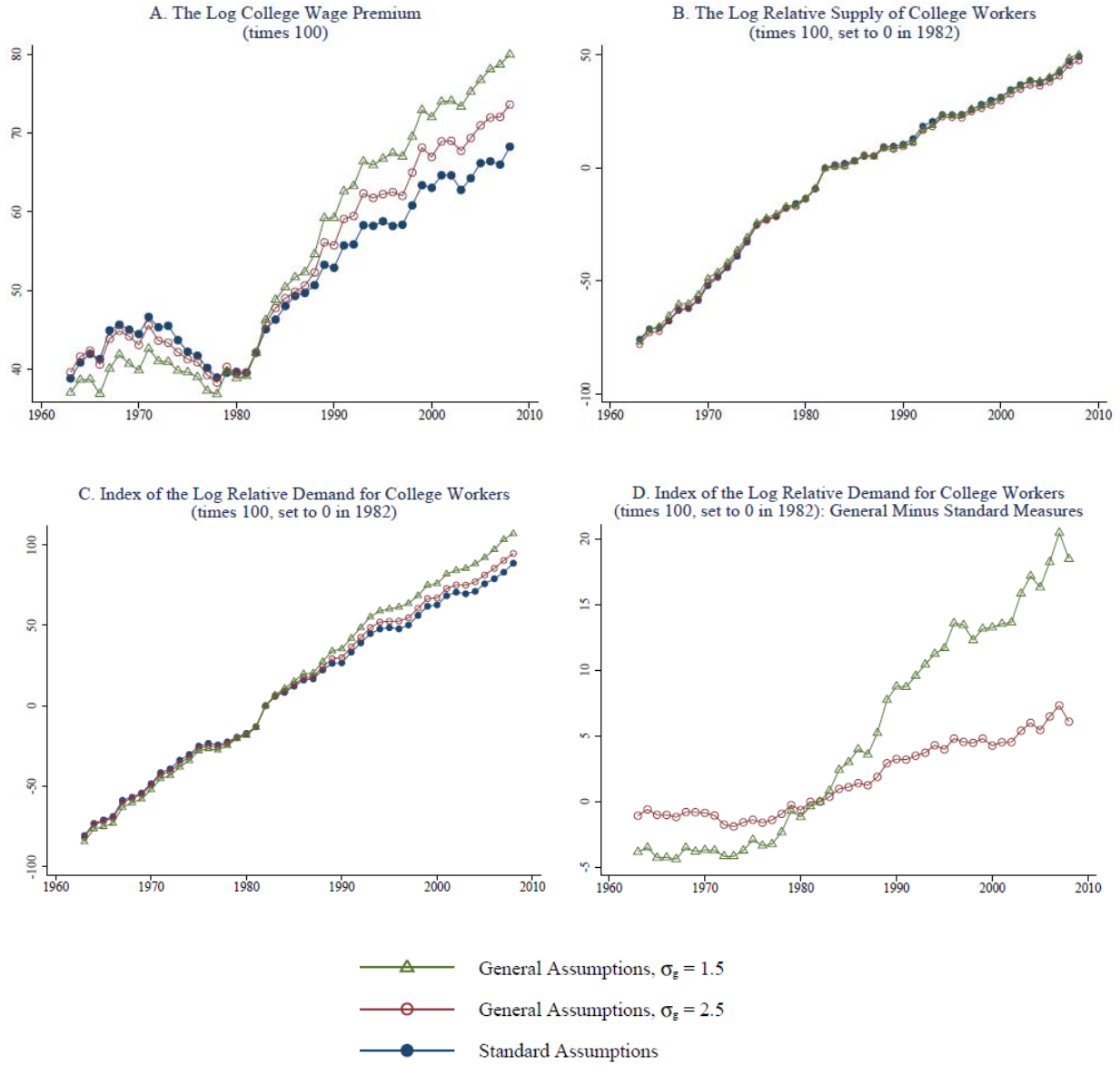
Figure 7: Measures of Log Wages and Supplies of Schooling Groups: Adjustments for Shifts in Supply and Demand within Schooling Groups

A. Log Wage Measures: Total Adjustment (multiplied by 100), Decomposed into Supply and Demand Effects

B. Log Supply Measures: Total Adjustment (multiplied by 100 and assuming $\sigma_g = 2.5$)



Figure 8: The Relative Wage of, Relative Supply of, and Relative Demand for College-Educated Workers, 1963-2008: Comparing Standard Measures and General Measures that Adjust for Shifts in Supply and Demand within Schooling Groups



I.F.2 Changes in the Relative Wage and Supply of College-Educated Labor

In this section, I investigate changes in measures of the relative wage and supply of college-educated labor: both the standard fixed-weight and adjusted measures described in equations (43) and (44) above. The key findings of this section are easily-summarized by inspecting Panels A and B of Figure 8. First, consider the relative supply series plotted in Panel B. The standard fixed-weight series (shifted by a constant) appears nearly identical to existing measures in the literature (e.g., Figure 2 of Acemoglu and Autor, 2011) and the long-run trend in its time series can, as originally noted by Katz and Murphy (1992), be broadly-characterized as a rapidly-increasing secular trend in the relative supply of the college-educated with a sharp slowdown in 1982. In fact, fitting this series to a simple linear trend with a trend break in 1982 is associated with an R-squared of 0.997.⁶¹ Panel B of Figure 8 also clearly-establishes that adjusting for supply and demand shifts within schooling groups does not greatly-alter the relative supply series; to illustrate this similarity, simple regressions of adjusted relative supply series with values of σ_g ranging from 1.5 to 4.5 on the standard fixed-weight series each feature an R-squared greater than 0.999.⁶²

Panel A of Figure 8 tells a different story for measures of the log college wage premium. Again, the standard fixed-weight series is very similar to existing measures in the literature (e.g., Figure 1 of Acemoglu and Autor, 2011): between 1963 and 1982, the fixed-weight log college wage premium fluctuates around 0.4; concurrent with the educational slowdown in Panel B, between 1982 and 1993, the series quickly increases to 0.58 during this 11-year period; beginning in 1993, however, the rapid increase of the fixed-weight log college wage premium slows considerably, increasing by 0.1 log points in the next 15 years to reach a value of 0.68 in 2008. In contrast with the log relative supply series, however, adjusted relative wage series differ significantly from the standard fixed-weight measure; between 1963 and 1982, changes over time in measures of the log college wage premium are relatively-similar; after 1982, however, growth in the adjusted log college wage premia begin to significantly-outpace growth in the standard fixed-weight log wage premium, implying that shifts in supply and demand within schooling groups depressed the standard measure of the college wage premium across

⁶¹Importantly, *fluctuations* in relative supply around long-run trends are strikingly-predictive of concurrent short-run movements in the college wage premium; here I simply establish the main long-run patterns of the relative log supply series. Card and DiNardo (2002) perform a similar exercise on observations from 1967 to 2000 and find an R-squared of 0.997. When I repeat the exercise using these earnings years, I find a similar R-squared value of 0.995.

⁶²All applications of the general model in this section use my preferred value of $\widehat{\sigma_e} = 5.025$ estimated in Section IV.

this period.

Taken together, these patterns suggest that, after adjusting for shifts in supply and demand within schooling groups, growth in the relative demand for college-educated labor after the early 1980s may be more-rapid than implied by estimates in the standard model. Before formally-investigating this hypothesis, I first briefly explore in more detail the sources of the adjustments to fixed-weight measures of log relative wages and supplies evident from Panels A and B of Figure 8: Figure 7 presents the entire time-series of these adjustments for particular parameter values, while Table 6 displays changes over time in these adjustments for a wider range of parameter values over a specific set of time periods: 1963-1982, 1982-1993, 1993-2008 and 1982-2008.⁶³

⁶³I describe in detail the rationale for choosing these particular time periods in the section on estimated changes in the relative demand for college-educated labor below.

TABLE 6
Annual Changes (times 100) in Log Wages and Log Supplies, 1963-2008:
Standard Fixed-Weight Measures and Adjustments for Shifts in Supply and Demand within Schooling Groups

	A. College Labor						B. High School Labor						C. College, Relative to High School Labor					
	1963-1982	1982-1993	1993-2008	1982-1993	1993-2008	1982-2008	1963-1982	1982-1993	1993-2008	1982-1993	1993-2008	1982-2008	1963-1982	1982-1993	1993-2008	1982-1993	1993-2008	1982-2008
Log Wage																		
Standard Fixed-Weight Log Wage Measure																		
Adjustments for Shifts Across Potential Experience Groups ($\sigma_e=5.025$):																		
Demand Shifts	-0.01	0.08	0.11	0.10			-0.01	0.03	0.04	0.03			0.00	0.05	0.08	0.08	0.06	
Supply Shifts	-0.01	0.04	-0.02	0.01			-0.01	0.01	-0.01	0.00			0.00	0.02	-0.01	0.00	0.00	
Adjustments for Shifts Across Gender:																		
$\sigma_g=1.5$																		
Demand Shifts	0.46	0.99	0.26	0.57			0.37	0.33	0.01	0.15			0.09	0.66	0.24	0.42		
Supply Shifts	0.09	-0.01	-0.07	-0.04			0.08	-0.01	0.01	0.00			0.01	0.01	-0.07	-0.04		
$\sigma_g=2.5$																		
Demand Shifts	0.10	0.45	0.08	0.23			0.13	0.17	0.02	0.08			-0.03	0.27	0.06	0.15		
Supply Shifts	0.04	0.02	-0.04	-0.01			0.04	0.00	0.01	0.00			-0.01	0.02	-0.04	-0.02		
$\sigma_g=3.5$																		
Demand Shifts	0.02	0.34	0.04	0.17			0.08	0.14	0.02	0.07			-0.06	0.20	0.03	0.10		
Supply Shifts	0.03	0.02	-0.03	-0.01			0.04	0.01	0.00	0.01			-0.01	0.02	-0.04	-0.01		
$\sigma_g=4.5$																		
Demand Shifts	-0.01	0.29	0.03	0.14			0.06	0.13	0.02	0.06			-0.07	0.17	0.01	0.08		
Supply Shifts	0.02	0.03	-0.03	-0.01			0.03	0.01	0.00	0.01			-0.01	0.02	-0.03	-0.01		
Sum: Total Adjustments																		
$\sigma_g=1.5$	0.53	1.11	0.29	0.63			0.44	0.37	0.05	0.18			0.09	0.74	0.24	0.45		
$\sigma_g=2.5$	0.11	0.59	0.14	0.33			0.16	0.22	0.05	0.12			-0.04	0.37	0.09	0.21		
$\sigma_g=3.5$	0.03	0.48	0.11	0.27			0.10	0.19	0.05	0.11			-0.07	0.29	0.06	0.16		
$\sigma_g=4.5$	0.00	0.44	0.09	0.24			0.08	0.18	0.05	0.1			-0.08	0.26	0.04	0.14		
Log Supply																		
Standard Fixed-Weight Log Supply Measure																		
Adjustments for Supply and Demand Shifts:																		
$\sigma_g=1.5$	3.06	1.71	1.61	1.65			-0.93	-0.16	-0.31	-0.25			3.99	1.87	1.92	1.90		
$\sigma_g=2.5$	-0.27	-0.14	0.15	0.03			-0.33	0.02	-0.03	-0.01			0.06	-0.16	0.18	0.04		
$\sigma_g=3.5$	-0.03	-0.27	0.00	-0.11			-0.15	-0.06	-0.03	-0.04			0.12	-0.21	0.03	-0.07		
$\sigma_g=4.5$	0.02	-0.29	-0.03	-0.14			-0.11	-0.07	-0.03	-0.04			0.13	-0.22	0.00	-0.10		
	0.05	-0.30	-0.04	-0.15			-0.09	-0.08	-0.02	-0.05			0.14	-0.22	-0.02	-0.10		

Adjustments to the Fixed-Weight Measure of Relative Supply The final four rows of Table 6 and Panel B of Figure 7 explore in detail, for a wide range of parameter values, the gap between the adjusted and fixed-weight measures of the log relative supply of college-educated labor outlined in equations (35) through (44) above. Confirming the evidence presented in Panel B of Figure 8, Panel C of Table 6 shows that change over time in relative supply is relatively unaffected by shifts in supply and demand within schooling groups: between 1963 and 1982, as the fixed-weight log relative supply of college-educated labor increased at nearly 4% annually, annual changes in the gap between adjusted and fixed-weight measures accounted for annual increases of between 0.06-0.14% overall, or between 1-3% of the total increase. Similarly, as the fixed-weight log relative supply of college-educated labor increased at the slower rate of roughly 1.9% annually from 1982 through 2008, adjustments account for between -6% and 2% of the overall increase in relative supply over the period.

What Decreased the (Fixed-Weight) College Wage Premium After 1982? A more-detailed examination of adjustments to wage measures is possible since, as shown in equations (35), (37) through (40), and (43), these adjustments can be decomposed into four components: shifts in both supply and demand within schooling groups across both the gender and potential experience dimensions. Inspection of Table 6 and Panel A of Figure 7 reveals that shifts in demand, not supply, within schooling groups have decreased the fixed-weight log college wage premium over the past three decades. For example, the rightmost column of Table 6 suggests that, during the 1982-2008 period, the sum of all shifts in supply accounted for roughly 10% of the (gross) changes in adjustments to the log fixed-weight college wage premium across all specifications. Two shifts in relative demand within schooling groups - across gender and across potential experience groups - appear to be driving the wage effects documented in Panel A of Figure 8.

Shifts in Demand Favoring Women Inspection of Panel C of Table 6 suggests that, after 1982, shifts in demand favoring women within both schooling groups played a primary role in decreasing the fixed-weight log college wage premium. Across specifications, these shifts in demand pushed down the college wage premium by between roughly 0.08-0.42% annually, or between 2-11% across the entire 1982-2008 period. Panel A of Figure 7 shows that these wage effects, within both groups, appear to have begun in the late 1970s, consistent with the evidence presented in Section V. Examination of equation (41) reveals that shifts in labor demand favoring female labor enter equation (43) as

employment-weighted effects: that is, gender-specific shifts in demand are weighted by gender-specific employment shares within each schooling group, displayed in Panel A of Table 4. Inspection of Table 4 shows that, within both schooling groups, women have supplied (on average, across the 1963-2008 earnings years) significantly less than half of the total usual hours worked: 35% among the college-educated and 41% among the high school educated. Two forces combine to push down the composition-adjusted college wage premium after 1982: first, since the female employment share is lower among the college-educated than among the high school-educated, even identical labor demand shifts favoring women will decrease the composition-adjusted log college wage premium; however, evidence presented in Section V suggests that shifts in labor demand favoring women are larger among the college-educated, further decreasing the composition-adjusted log college wage premium after 1982. Inspection of Panel C of Table 6 shows that these effects are sensitive to the elasticity of substitution between gender, σ_g . This is consistent with observed changes in the relative wage and supply of female labor presented in Table 5 and Figure 6: after 1982, although changes in the gender wage gap are similar within both college- and high school-educated workers, growth the relative supply of female labor within the college-educated far-outstrips relative supply growth within the high school-educated. As a result, if women and men are less-substitutable, larger shifts in relative demand within college-educated workers are required to reconcile observed patterns of relative wages and supplies.⁶⁴

Shifts in Demand Favoring the Experienced Among Women Row 2 of Table 6 shows that shifts in relative demand across potential experience groups combine to decrease the college wage premium by roughly 0.06% annually, or by 2% across the entire 1982-2008 period. This effect is primarily the result of shifts in demand favoring workers with more potential experience among college-educated women beginning in the late 1970s, documented in Panel A of Table 1 and Panel C of Figure 4. To see this, first note that, similar to shifts in demand across gender, shifts in demand across potential experience groups enter equation (43) as employment-weighted effects: equations (28) and (39) show that shifts in demand towards demographic groups that are less-represented within a given schooling-by-gender group will tend to increase the wage effect ξ_{st}^{De} over time. Panel A of Table 4 reveals that, on average across the 1963-2008 earnings years, roughly half of the total usual hours worked among the high school-educated are supplied by workers with 20+ years of experience: 47%

⁶⁴ As noted in Section V, the reason this pattern is not observed in the relative demand *indices* \widehat{D}_{st} displayed in Table 5 and Figure 6 is that these indices are weighted by $(\widehat{\sigma}_g - 1)$.

among women and 45% among men. As a result, shifts in labor demand towards more-experienced workers will have relatively-little effect on the high school-educated wage, as can be seen in Panel B of Table 6 and in Panel A.2 of Figure 7. However, Table 4 shows that only 35% of the average total usual hours worked among college-educated women are supplied by workers with 20+ years of experience; as a result, shifts in labor demand favoring these workers, and documented in Section IV, will push down the (composition-adjusted) college-educated wage, and hence the (composition-adjusted) college wage premium.

I.F.3 Estimated Changes in the Relative Demand for College-Educated Labor

The key question of this paper is how shifts in supply and demand within schooling groups affect the measurement of the relative demand for college-educated labor; evidence presented in the previous section suggests that, across a wide range of values for σ_g , shifts in relative demand favoring women, and favoring workers with more potential experience among women, left standard fixed-weight measures of log relative supply relatively unchanged, but depressed the fixed-weight measure of the log college wage premium after 1982 for a total decrease of between 4-12%. These results suggest that, after adjusting for shifts within schooling groups, relative demand for college-educated workers may have increased more-quickly after 1982 than as measured in the standard model.

The Late 1970s: Adjusted Growth Begins to Outpace the Standard Measure Panels C and D of Figure 8 confirm this suggestion. These panels present - for a value of the elasticity of substitution between schooling groups of $\sigma_s = 1.5$ and for values of σ_g equal to 1.5 and 2.5 - change over time in the standard index of the relative demand for college-educated workers \widehat{D}_t , as defined in equation (8) and expressed in equation (45), and computed either imposing "standard" assumptions (within each schooling group, demand is assumed to be held fixed and workers are assumed to be perfect substitutes) which force lines 2 and 3 of equation (45) to equal zero, or under the general case of the model, which relaxes standard assumptions and allows lines 2 and 3 of equation (45) to vary.

First, consider the standard measure of \widehat{D}_t plotted in Panel C of Figure 8. Both long- and short-run fluctuations in this series appear very similar to existing series in the literature (e.g., Panel D of Figure 4 in Katz and Murphy, 1992), and reflect the notion that, across the 1963-2008 earnings years, observed trends in the relative wage and supply of college-educated labor, displayed in Panels A and B of Figure 8, are potentially-reconciled by a long-run increasing trend in the relative demand for college-

educated workers.⁶⁵ However, beginning in the late 1970s, concurrent with the within-schooling group demand shifts discussed above, the general measure of \widehat{D}_t begins to increase more-rapidly than the standard measure. Panel D of Figure 8 suggests that, for $\sigma_s = 1.5$ and values of σ_g between 1.5 and 2.5, the general measure of relative demand for college-educated workers increased by roughly 8-21% more than its standard counterpart since 1975.

⁶⁵Evidence suggests that this long-run trend is important across the last century of U.S. history (Goldin and Katz, 2008; 2009). Panel B of Figure 4 of Katz and Murphy (1992) suggests that, for a value of σ_s of 1.41, \widehat{D}_t increased by slightly more than 100% from 1963 to 1987. The same calculation using standard assumptions and a value of $\sigma_s = 1.5$ from Panel C of Figure 8 of the current paper implies that \widehat{D}_t increased by approximately 98% over these years.

TABLE 7

Annual Changes (times 100) in the Index of Relative Demand for College-Educated Labor, 1963-2008:
The Standard Measure and Adjustments for Shifts in Supply and Demand within Schooling Groups

The Standard Measure	A. $\sigma_s = 1.29$				B. $\sigma_s = 1.5$				C. $\sigma_s = 1.84$			
	1963- 1982	1982- 1993	1993- 2008	1982- 2008	1963- 1982	1982- 1993	1993- 2008	1982- 2008	1963- 1982	1982- 1993	1993- 2008	1982- 2008
Decomposition into Wage and Supply Components:												
Fixed-Weight College Wage Premium	0.22	1.90	0.86	1.30	0.26	2.21	1.00	1.51	0.32	2.71	1.22	1.85
Fixed-Weight Relative Supply of the College-Educated	3.99	1.87	1.92	1.90	3.99	1.87	1.92	1.90	3.99	1.87	1.92	1.90
Sum: Total Standard Measure	4.21	3.77	2.78	3.20	4.25	4.08	2.92	3.41	4.31	4.58	3.14	3.75
Adjustments for Shifts in Supply and Demand within College and High School Labor												
Adjustments to the College Wage Premium Component:												
Shifts Across Potential Experience Groups ($\sigma_e=5.025$):												
Demand Shifts	-0.01	0.07	0.10	0.08	-0.01	0.08	0.11	0.10	-0.01	0.09	0.14	0.12
Supply Shifts	0.00	0.03	-0.01	0.01	0.00	0.04	-0.01	0.01	0.00	0.05	-0.02	0.01
Shifts Across Gender:												
$\sigma_g=1.5$												
Demand Shifts	0.12	0.85	0.32	0.54	0.14	0.98	0.37	0.63	0.17	1.21	0.45	0.77
Supply Shifts	0.01	0.01	-0.09	-0.05	0.01	0.01	-0.11	-0.06	0.01	0.01	-0.13	-0.07
$\sigma_g=2.5$												
Demand Shifts	-0.04	0.35	0.08	0.20	-0.05	0.41	0.09	0.23	-0.06	0.50	0.12	0.28
Supply Shifts	-0.01	0.02	-0.05	-0.02	-0.01	0.03	-0.06	-0.03	-0.01	0.03	-0.08	-0.03
$\sigma_g=3.5$												
Demand Shifts	-0.07	0.26	0.03	0.13	-0.09	0.30	0.04	0.15	-0.11	0.36	0.05	0.18
Supply Shifts	-0.01	0.02	-0.05	-0.02	-0.01	0.03	-0.05	-0.02	-0.01	0.03	-0.07	-0.02
$\sigma_g=4.5$												
Demand Shifts	-0.09	0.21	0.01	0.10	-0.10	0.25	0.02	0.11	-0.13	0.30	0.02	0.14
Supply Shifts	-0.01	0.03	-0.04	-0.01	-0.01	0.03	-0.05	-0.02	-0.02	0.04	-0.06	-0.02
Adjustments to the Relative Supply Component:												
$\sigma_g=1.5$												
Demand Shifts	0.06	-0.16	0.18	0.04	0.06	-0.16	0.18	0.04	0.06	-0.16	0.18	0.04
Supply Shifts	0.12	-0.21	0.03	-0.07	0.12	-0.21	0.03	-0.07	0.12	-0.21	0.03	-0.07
$\sigma_g=2.5$												
Demand Shifts	0.13	-0.22	0.00	-0.10	0.13	-0.22	0.00	-0.10	0.13	-0.22	0.00	-0.10
Supply Shifts	0.14	-0.22	-0.02	-0.10	0.14	-0.22	-0.02	-0.10	0.14	-0.22	-0.02	-0.10
$\sigma_g=3.5$												
Demand Shifts	0.18	0.79	0.49	0.62	0.20	0.95	0.54	0.71	0.23	1.20	0.62	0.86
Supply Shifts	0.06	0.26	0.14	0.19	0.06	0.34	0.16	0.23	0.04	0.46	0.19	0.30
$\sigma_g=4.5$												
Demand Shifts	0.04	0.16	0.07	0.11	0.03	0.22	0.08	0.14	0.00	0.32	0.10	0.19
Supply Shifts	0.03	0.11	0.04	0.07	0.02	0.17	0.05	0.10	-0.01	0.25	0.06	0.14
Sum: Total Adjustments to the Standard Measure												
$\sigma_g=1.5$												
Demand Shifts	0.18	0.79	0.49	0.62	0.20	0.95	0.54	0.71	0.23	1.20	0.62	0.86
Supply Shifts	0.06	0.26	0.14	0.19	0.06	0.34	0.16	0.23	0.04	0.46	0.19	0.30
$\sigma_g=2.5$												
Demand Shifts	0.13	-0.22	0.00	-0.10	0.13	-0.22	0.00	-0.10	0.13	-0.22	0.00	-0.10
Supply Shifts	0.14	-0.22	-0.02	-0.10	0.14	-0.22	-0.02	-0.10	0.14	-0.22	-0.02	-0.10
$\sigma_g=3.5$												
Demand Shifts	0.18	0.79	0.49	0.62	0.20	0.95	0.54	0.71	0.23	1.20	0.62	0.86
Supply Shifts	0.06	0.26	0.14	0.19	0.06	0.34	0.16	0.23	0.04	0.46	0.19	0.30
$\sigma_g=4.5$												
Demand Shifts	0.04	0.16	0.07	0.11	0.03	0.22	0.08	0.14	0.00	0.32	0.10	0.19
Supply Shifts	0.03	0.11	0.04	0.07	0.02	0.17	0.05	0.10	-0.01	0.25	0.06	0.14

Estimated Changes in the Relative Demand for College-Educated Labor While Figure 8 allows for inspection of the full time series of \widehat{D}_t , Table 7 presents evidence on changes over time in \widehat{D}_t that are (i) computed for a wider range of values for both σ_s and σ_g , and are (ii) decomposed into the components outlined in equation (45) above, so that the precise driving forces of the measured differences between general and standard measures can be illuminated.

Before turning to the figures in the table, I first discuss the periods I consider. I examine changes in \widehat{D}_t across the periods 1963-1982, 1982-1993 and 1993-2008. I choose these periods for two reasons. First, they divide the 45 years between 1963-2008 into their most-meaningful components in terms of changes in the relative wage and supply of college-educated labor (as seen in Panels A and B of Figure 8): from 1963-1982, the log college wage premium fluctuated around 0.4 while the relative supply of the college-educated increased quickly; in 1982, the college wage premium increased dramatically, concurrent with the well-documented slowdown in the educational attainment of the labor force, and finally in 1993 increase in the (fixed-weight) college wage premium began to increase less-dramatically than during the previous 11 years. A second reason is that, during the 1976-1981 period, the standard model performs relatively poorly in predicting the time series of the college wage premium, potentially as a result of non-market factors, such as the union-negotiated wage settlements of the 1970s, that complicate the measurement of relative demand growth across periods that include this range of earnings years (Goldin and Katz, 2009). Thus, in order to meaningfully-compare increase in the growth of relative demand surrounding these years, for the purposes of Table 7, I choose to combine the years 1963-1982 into one continuous period; inspection of Panel C of Figure 8 shows that, for the value $\sigma_s = 1.5$, this empirical choice is relatively-similar to considering changes in \widehat{D}_t over the alternative period 1963-1975.

To investigate this difference in more detail, Table 7 presents changes over time in \widehat{D}_t under both standard and general assumptions for a wide range of parameter values. First, Panels A, B and C employ values of the elasticity of substitution between schooling groups σ_s of 1.29, 1.5 and 1.84, respectively, that are implied by estimates in columns 6, 10, and 12 of Table 8 (discussed in the next section), respectively, and which together span the range of values of σ_s implied by the preferred estimates of Goldin and Katz (2009; 1.4, 1.64 and 1.84 examined in their Table 1) and Acemoglu and Autor (2011; 1.55, 1.78 and 1.80 estimated in columns 3, 4, and 5 of their Table 8). Second, rows of the table display a range of values for σ_g of between 1.5 and 4.5 that, as discussed in Section V, is guided by existing estimates in the literature (Acemoglu and Autor, 2004; Blau and Kahn, 2004; Layard,

1982; Lewis, 1985; Weinberg, 2000). When displaying results associated with the general model, I take 1.5 as my preferred estimate of σ_s , since this value is implied by the standard trend-break model of the college wage premium in column 10 of Table 8, and I take 2.5 as my preferred estimate of σ_g , since this value is closest to the median value of the preferred estimates of the papers referenced in the previous sentence. Rows of Table 7 explicitly follow the form of equation (45): the first three rows of Table 7 present the standard measure of growth in the relative demand for college-educated workers from line 1 of the RHS of equation (45), while the following rows present estimates of the adjustments contained in lines 2 and 3 of equation (45).

Inspection of Table 7 yields four main take-aways. First, across all but one of the 48 displayed specifications and time periods, growth in the relative demand for college-educated workers is more-rapid after adjusting for shifts in supply and demand within schooling groups.⁶⁶ Second, increases are concentrated after 1982: while estimates vary significantly across specification, roughly 10-60% of the gap in the annual growth rates of the (standard measure of) relative demand for college-educated workers between the 1963-1982 and 1982-2008 periods can be closed by adjusting for shifts in supply and demand within schooling groups. Across the values of σ_s displayed in row 3 of Table 7, the standard measure of the annual rate of relative demand growth decreased by between roughly 0.55% ($\sigma_s = 1.84$) and 1.01% ($\sigma_s = 1.29$) between these periods. For my preferred estimate of $\sigma_s = 1.5$ (estimation of which I discuss in the next section), adjusting for shifts in supply and demand within schooling groups decreases this measured gap by between roughly 0.08% ($\sigma_g = 4.5$) and 0.51% ($\sigma_g = 1.5$) annually, or between roughly 10-61% as a fraction of the gap. Similarly, for my preferred estimate of $\sigma_g = 2.5$, this measured gap decreases by between roughly 0.13% ($\sigma_s = 1.29$) and 0.26% ($\sigma_s = 1.84$) annually, or between roughly 12-47% as a fraction of the gap.⁶⁷

Third, and as suggested by the wage effect decompositions presented in Table 6 and discussed in the previous section, these effects are mainly driven by shifts in relative demand within schooling groups, and are sensitive to the value of σ_g . For example, figures in Panel B of Table 7 imply that, for my

⁶⁶The lone exception is the adjustment of -0.01 across the 1963-1982 period implied by the values $\sigma_s = 1.84$ and $\sigma_g = 4.5$ and displayed in Panel C of Table 7.

⁶⁷Note that the estimates presented in row 3 of Table 7 are close to existing estimates in the literature. Goldin and Katz (2009, Table 1) suggest that, for their preferred value of $\sigma_s = 1.64$, annual growth in \widehat{D}_t was approximately 3.73 from 1960-1980 and 3.48 from 1980-2005. Estimates from Panel C of Table 2 of Autor et al. (1998) associated with their preferred value of $\sigma_s = 1.4$ suggest annual growth in \widehat{D}_t of 3.74 for the 1960-1980 period and 4.02 across the 1980-1996 earnings years. My estimates of annual growth in \widehat{D}_t for the preferred value of $\sigma_s = 1.5$ from Panel B of Table 7 are 4.25 from 1963-1982, 4.08 from 1982-1993 and 3.41 from 1982-2008.

preferred estimate of $\sigma_s = 1.5$, demand shifts account for between 62-87% of the (gross) adjustment to changes in the relative demand for the college-educated across the 1982-2008 period. Across these earnings years and for this value of σ_s , adjustments to the annual growth rate of \widehat{D}_t due to gender-specific demand shifts vary from roughly 0.11-0.63% annually, depending on this value. Finally, because adjustments to the growth rate of \widehat{D}_t are generally largest during the 1982-1993 period, as shifts in labor demand favoring women were most-rapid (as measured in Section V of the paper), the well-documented decrease in the growth rate of \widehat{D}_t after 1993 (Autor et al., 2008; Goldin and Katz, 2009) remains across all specifications. For example, in the case of my preferred specification of $\sigma_s = 1.5$ and $\sigma_g = 2.5$ plotted in Panel C of Figure 8 and displayed in Panel B of Table 7, adjusted annual growth in the relative demand for college-educated workers decreased from roughly 4.42% during the 1982-1993 period to roughly 3.07% across the 1993-2008 earnings years.

TABLE 8
Estimating Models of Changes in the Log College Wage Premium: Comparing the Standard Model and
Generalized Models that Adjust for Shifts in Supply and Demand within Schooling Groups

	A. Standard Model of the Fixed- Weight Log College Wage Premium				B. Generalized Model of the Adjusted Log College Wage Premium ($\sigma_{\varepsilon} = 1.5$)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fixed-Weight Log Relative Supply Measure (Standard Model) / Adjusted Log Relative Supply Measure (Generalized Model)	-0.430** [0.045]	-0.582** [0.072]	-0.643** [0.071]	-0.564** [0.092]	-0.573** [0.055]	-0.777** [0.078]	-0.849** [0.087]	-0.535** [0.124]
Time	0.018** [0.001]	0.024** [0.002]	0.029** [0.003]	0.021** [0.006]	0.026** [0.002]	0.033** [0.003]	0.039** [0.004]	0.016 [0.008]
post-1992		0.195** [0.042]				0.264** [0.045]		
Time * post-1992		-0.006** [0.001]				-0.008** [0.001]		
Time ² / 100			-0.010** [0.003]	0.016 [0.016]			-0.013** [0.003]	0.054* [0.022]
Time ³ / 1000				-0.003 [0.002]				-0.008** [0.003]
Constant	-0.305** [0.073]	-0.556** [0.119]	-0.671** [0.120]	-0.523** [0.160]	-0.796** [0.094]	-1.152** [0.136]	-1.291** [0.154]	-0.701** [0.228]
Observations	46	46	46	46	46	46	46	46
R-Squared	0.943	0.958	0.948	0.952	0.969	0.981	0.972	0.978
R-Squared: Predicted vs. Fixed-Weight Log College Wage Premium, Over Displayed Period								
1963-2008	0.943	0.958	0.948	0.952	0.921	0.951	0.929	0.945
1963-1975	0.691	0.498	0.712	0.410	0.797	0.655	0.762	0.330
1982-2008	0.955	0.973	0.962	0.967	0.933	0.971	0.949	0.960
	C. Generalized Model of the Adjusted Log College Wage Premium ($\sigma_{\varepsilon} = 2.5$)				D. Generalized Model of the Adjusted Log College Wage Premium ($\sigma_{\varepsilon} = 3.5$)			
	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
Adjusted Log Relative Supply Measure	-0.495** [0.045]	-0.668** [0.070]	-0.733** [0.076]	-0.544** [0.108]	-0.482** [0.043]	-0.647** [0.068]	-0.712** [0.075]	-0.545** [0.105]
Time	0.021** [0.001]	0.027** [0.002]	0.033** [0.003]	0.018* [0.007]	0.021** [0.001]	0.026** [0.002]	0.032** [0.003]	0.018* [0.007]
post-1992		0.231** [0.043]				0.223** [0.043]		
Time * post-1992		-0.007** [0.001]				-0.007** [0.001]		
Time ² / 100			-0.012** [0.003]	0.035 [0.020]			-0.012** [0.003]	0.031 [0.020]
Time ³ / 1000				-0.006* [0.002]				-0.005* [0.002]
Constant	-0.609** [0.078]	-0.914** [0.122]	-1.041** [0.137]	-0.679** [0.200]	-0.575** [0.075]	-0.866** [0.120]	-0.993** [0.134]	-0.672** [0.196]
Observations	46	46	46	46	46	46	46	46
R-Squared	0.958	0.972	0.962	0.968	0.954	0.969	0.959	0.964
R-Squared: Predicted vs. Fixed-Weight Log College Wage Premium, Over Displayed Period								
1963-2008	0.935	0.956	0.941	0.950	0.937	0.957	0.943	0.951
1963-1975	0.789	0.629	0.763	0.394	0.784	0.620	0.760	0.405
1982-2008	0.945	0.973	0.958	0.965	0.947	0.972	0.959	0.966

I.F.4 Estimating a (Katz Murphy) Model of the College Wage Premium

The previous section presented estimates of growth in the relative demand for college-educated workers for a broad range of parameter values. This is likely the most-complete approach, given the well-known problems in estimating elasticities such as σ_s with time-series regressions such as equation (43), as noted in the original article of Katz and Murphy (1992). However, it is still instructive to examine the effects of shifts in supply and demand within schooling groups on estimates of the parameter σ_s . In Table 8, I follow the classic approach of Katz and Murphy in estimating σ_s using equation (43): in Panel A, I examine estimates of σ_s implied from versions of equation (43) in which I impose the assumptions of the standard model, which force adjustments to fixed-weighted measures of log relative wages and supplies in the equation to equal zero, while in Panels B, C, and D, I examine estimates implied from relaxing these assumptions, which allows the adjustment terms in equations (43) and (44) to be non-zero (for the values $\sigma_g = 1.5, 2.5$ and 3.5 , respectively).

Estimates of σ_s derived from Panel A are very close to those estimated in existing literature: estimates of 2.33, 1.72, 1.56 and 1.77 implied by columns 1, 2, 3 and 4 of Panel A, respectively, are statistically-indistinguishable at the 5% level from estimates of 2.95, 1.55, 1.78 and 1.80 implied by identical regressions displayed in columns 2, 3, 4, and 5, respectively, of Table 8 in Acemoglu and Autor (2011).⁶⁸ Inspection of Table 8 clarifies three basic differences between the standard specifications estimated in Panel A and estimates of the general versions of equation (43) presented in Panels B through D. First, since adjusting for shifts in supply and demand within schooling groups, as shown in the previous section, tends to increase the measured rate of growth of the relative demand for college-educated workers across all time periods, it is unsurprising that the time trends estimated in Panels B, C and D are steeper than their counterparts in Panel A. The measured coefficients on the time variable in the standard model in columns 1, 2 and 3 of Panel A of Table 8, for example, are each lower than the associated coefficients in the analogous columns of Panels B-D (though only the measured gaps in these coefficients implied by comparing Panels A and B are statistically-distinguishable at the

⁶⁸Estimates in Acemoglu and Autor use the same source data and nearly-identical empirical methods. However, one possible explanation for the small measured gap between estimates found in their Table 8 and estimates in Panel A of Table 8 in the current paper is that employment weights in their paper (analogous to the employment weights $\widehat{\lambda}_{sge}$ employed in this paper, and displayed in Table 4) appear to be averaged across the 1963-1987 earnings years (to coincide with column 1 of their Table 8, which is meant to replicate the original findings of Katz and Murphy, 1992), whereas my employment weights are, as is standard, averaged across all available earnings years, 1963-2008. Re-computing estimates in Panel A of Table 8 of the current paper with values of $\widehat{\lambda}_{sge}$ averaged over the 1963-1987 period (not displayed) yields estimates considerably closer to those of Acemoglu and Autor.

5% level).

A second, and complementary, result is that the generalized models in Panels B-D are, in general, associated with smaller estimated values for σ_s than their counterparts in Panel A, implying that college- and high school-equivalent workers (weighted by time-varying parameters as made explicit in equations 2 and 3) are less-substitutable than in the standard model. In 9 out of 12 comparisons between estimates in columns 5-16 and the corresponding standard estimates in columns 1-4, the values of σ_s implied by the generalized model are smaller than those implied by the standard model (though in none of these cases can the null hypothesis that the difference equals zero be rejected at the 5% level).⁶⁹ For example, my preferred model for the log college wage premium in column 10 - a standard specification that yields a value of $\sigma_s = 1.64$ in the benchmark work of Goldin and Katz (2009) - yields the estimated value of $\sigma_s = 1.5$. These findings are intuitive: in the general model, a larger portion of the post-1982 increase in the (adjusted) log college wage premium is attributable to increases in the (adjusted) relative demand for college-educated workers; slightly less of this increase is, then, attributable to the post-1982 slowdown in the educational attainment of the labor force, implying a smaller elasticity of substitution.

Third, and finally, comparisons of Panels A and B-D are necessarily comparisons of "apples to oranges:" while the object of interest in Panel A is the fixed-weight log college wage premium, the dependent variables in Panels B-D are different quantities: *adjusted* wage gaps. As a result, comparisons of the simple displayed OLS R-squared statistics - generally higher for the generalized model - are not useful in comparing the predictive power of the standard versus generalized specifications. In the final 3 rows of each panel, I therefore present R-squared statistics associated with simple regressions of the fixed-weight log college wage premium on its predicted value. Comparisons of these three rows between Panel A and Panels B-D suggest that the fit of the general model is relatively-similar to that of the standard model. A leading example is the standard trend-break specification in columns 2, 6, 10 and 14 of the table. Comparisons of the R-squared statistic across the entire 1963-2008 period suggests that the fit of the standard specification is slightly better overall: R-squared values are 0.958 in column 2, but range from between 0.951 and 0.957 for the general models. A slightly-different story is found by extracting the years 1976-1981, during which the fit of both models is poor (for reasons discussed above and in, e.g., Goldin and Katz, 2009). Across the 1963-1975 earnings years, during which

⁶⁹The three exceptions are each associated with modeling changes in relative demand with a 3rd-order polynomial: columns 8, 12 and 16.

the fixed-weight log college wage premium fluctuated around 0.4, the fit of each general specification is slightly better than that of the standard specification: the R-squared value for this period in column 2 is 0.498 but ranges from 0.620 to 0.655 for the analogous general specifications. Across the 1982-2008 earnings years, during which the fixed-weight log college wage premium increased dramatically, the fit of the standard and general specifications under consideration are very similar: R-squared estimates associated with this period in columns 2, 6, 10 and 14 each roughly equal 0.972. Taken together, these rows suggest that the fit of the standard and general models are roughly equivalent across the 1963-2008 earnings years.

I.G Conclusion

As measured in the standard two schooling group supply and demand model of the U.S. labor market, growth in the relative demand for college-educated labor underwent a puzzling slowdown during the computer revolution. While this slowdown is generally hypothesized to be driven by changes in technology that depressed demand for the routine tasks performed by middle-skilled workers - including some college-educated workers - in this paper I investigated the extent to which supply and demand shifts within schooling groups may have affected measurement of the relative demand for college-educated labor from 1963 to 2008. I find that, beginning in the late 1970s, shifts in labor demand became more-favorable to women relative to men and, among women, favored those with more (potential) labor market experience. A significant part of the slowdown after 1982 in the growth rate of the relative demand for college-educated labor can be explained by these demand shifts, but the extent of this effect varies considerably across specifications. Moreover, evidence presented in this paper is complementary to existing work on how computers changed the labor market: in particular, estimates suggest that the relative demand for college-educated workers may have decreased less than previously-thought in the years after the introduction of the IBM-PC in 1981, but appear to confirm the economic importance of the technological shifts responsible for the sharp slowdown in the relative demand for college-educated labor since the mid-1990s.

II Did Relative Labor Demand Shift Towards Less-Educated Workers in Brazil, 1995-2011?

II.A Introduction

In a sample of 16 countries that together comprise over 97% of the total Latin American population,⁷⁰ composition-adjusted measures of tertiary/secondary and secondary/less-than-secondary wage differentials decreased in 16 and 12 countries, respectively, by a (national) annual average of approximately 2.8% and 1%, respectively, during the 2000s (Gasparini et al., 2011; Tables 3.1 and 3.2).⁷¹ A recent literature directly-applies standard supply and demand frameworks, originally developed to analyze schooling differentials in the U.S. labor market (Card and Lemieux, 2001; Goldin and Katz, 2008; 2009; Katz and Murphy, 1992), to examine the sources of these striking decreases (Cruces et al., 2012; Galiani, 2009; Gasparini et al., 2011; Manacorda et al., 2010).⁷² A relatively-common pattern emerges: while many countries in Latin America feature smooth, long-run increases in the relative supply of more-educated workers,⁷³ schooling premia underwent sharp negative trend-breaks during the 2000s. One interpretation of these time series is that the sudden and roughly decade-long world-wide increase in commodity prices beginning in the early 2000s (Baffles and Haniotis, 2010; Helbling et al., 2008; Yueh, 2013) increased the relative demand for jobs performed by less-educated workers in commodity-exporting Latin American countries, driving down schooling premia during this decade (Gasparini et al., 2011).

In this paper, I first confirm that, among male wage and salary workers in Brazil, a simple schooling premium measure - the ratio of the composition-adjusted mean wage of men with upper-secondary,

⁷⁰ Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Uruguay and Venezuela.

⁷¹ Additional evidence on the time-series of schooling wage differentials in Latin America in the 2000s is presented in Lustig et al. (2011; 2012).

⁷² The Katz and Murphy (1992) supply and demand model (in Blom and Velez, 2004; Gallego, 2012; Montes Rojas, 2006) and Juhn Murphy and Pierce (1993) decomposition framework (in Sotomayor, 2004) have also been taken, with little alteration, to yearly datasets from Brazil and other Latin American countries to analyze earlier periods.

⁷³ For the case of Brazil, Plank (1987; Figure 2) shows that long-run, increasing secular trends in enrollment rates in nearly every state in Brazil began in earnest during the 1950s. These long-run secular trends appear to be associated with a combination of long-run increases in the demand for educated workers (Blom et al., 2001), as well as institutional reforms: the Brazilian Constitution of 1934 defined education as a basic right (Rodriguez et al., 2008), while the Constitution of 1946 established basic financial guarantees for public education (Plank, 1987).

versus less than upper-secondary, education - is stable at roughly 2.5 from 1995-2002, then compresses to 1.9 by 2011, consistent with the simple education story above. However, a standard two schooling group formalization of this story appears inconsistent with observed time series of relative wages and supplies, potentially suggesting that shifts in labor demand were heterogeneous *within* these broad education groups. This analysis implies that, in the Brazilian case, relevant shifts in demand - whatever their source - are not easily-characterized along the schooling dimension alone.

I begin by documenting the dramatic decrease in real wage inequality among males in Brazil across the 1995-2011 earnings years, and I show that these years can be meaningfully-decomposed into two distinct periods. First, from 1995 to 2002, most men experienced real wage losses as output in the economy stagnated. However, while wage decreases were relatively-similar for most men, the extreme left tail of the wage distribution compressed as those workers below the 20th percentile of the wage distribution saw real wage gains. Second, from 2002 to 2011, there was a dramatic compression of the log wage distribution that was monotonically-decreasing and near-linear in the percentile of the distribution as per-capita GDP increased by roughly 25%. Overall, from 1995-2011, the gap between the 90th and 10th percentiles of the (raw) wage distribution closed by roughly 4.6% annually: roughly 1.1% per year faster than the analogous gap among white males during the "Great Compression" of the U.S. weekly wage distribution from 1940-1950 (Goldin and Margo, 1992). Movements in composition-adjusted relative wages between men grouped by their level of schooling are also qualitatively different between these two periods. From 1995-2002, the real wages of men along all points in the schooling distribution fell together; during the boom years from 2002-2011, by contrast, wage increases were monotonically-decreasing in level of schooling: composition-adjusted real wages of workers with 0, 1-4, 5-8, 9-11 and 12+ years of schooling increased by roughly 54%, 24%, 16%, 1% and -12%, respectively, across this period.

These patterns, together with smooth increases in the relative supply of more-educated workers across the entire 1995-2011 period, appear to suggest a simple supply and demand story in which the relative demand for more-educated workers slowed in 2002. However, I show that a simple two schooling group model of the upper-secondary/non-upper-secondary wage premium, intended to match the theoretical choices of standard models of schooling premia used to analyze the decline of schooling premia in Brazil and other Latin American countries in the 2000s (Cruces et al., 2012; Galiani, 2009; Gasparini et al., 2011; Manacorda et al., 2010), appears to be inconsistent with the data. Specifically, the data reject the model's predictions that the time-series of schooling premia should be similar for

workers with different levels of potential experience, and that experience premia should be relatively constant within schooling groups. I find that these discrepancies between the model and data are unlikely to be driven by differential shifts in supply within schooling groups.

Finally, as an alternative to models that group workers solely by their level of schooling, I consider a model that groups workers not only by their level of schooling, but also by their level of labor market experience. I compare three types of workers: "low-skill" men are defined as those with neither advanced schooling or experience (0-8 years of schooling and 0-7 years of experience), "high-skill" men are defined as those with both advanced schooling and experience (9+ years of schooling and 18-35 years of experience), and "middle-skill" men are defined as those with one, but not both of those characteristics. Flexible application of this model to the 1995-2011 earnings years in Brazil implies that annual growth in the relative demand for high-, relative to middle-skill workers was roughly 4-12% lower from 2002-2011 than during the 1995-2002 period, while annual growth in the relative demand for middle-, relative to low-skill workers was roughly 6-9% faster. These results, however, must be interpreted as suggestive: I cannot rule out the notion that either changes in cohort quality or model mis-specification affect the estimation of relative demand shifts.

The remainder of the paper is organized as follows. In Section II, I outline the data and basic empirical procedures employed in the rest of the paper. In Section III, I document the basic patterns of both raw wage inequality, as well as relative wages and relative supplies of workers of different demographic groups, across the 1976-2011 earnings years. In Section IV, I evaluate different models for interpreting these patterns. Section V concludes.

II.B Data and Empirical Measures

In this section, I describe the data, basic processing procedures, and the empirical measures of wages and labor supply that I use in the paper. My approach is to use procedures that match, as closely as possible, the empirical choices of Katz and Murphy (1992). Although recent literature on U.S. relative wages (e.g., Autor et al., 2008) has modified these choices - perhaps most-notably, by using mean log wages, rather than log mean wages as wage measures - the bulk of the recent literature on Brazilian and Latin American relative wages (Blom and Velez, 2004; Cruces et al., 2012; Galiani, 2009; Gallego, 2012; Gasparini et al., 2011; Manacorda et al., 2010; Montes Rojas, 2006) adopts the original empirical methodology of Katz and Murphy. As a result, I follow the approach of these papers.

II.B.1 Data

The Pesquisa Nacional por Amostra de Domicílios (PNAD) is the longest-running yearly household survey in Brazil. Although the Brazilian Institute of Geography and Statistics began collecting the PNAD in 1967, nationally-representative samples began in 1976, and have been collected in each subsequent year, except for 1994 and years during which the Household Census was collected (1980, 1991 and 2000). The PNAD contains the worker-level information I use to construct yearly measures of wages and supplies of workers of different demographic groups.

First, the PNAD provides information on employment: the survey asks, in reference to a fixed "reference week," whether the respondent was employed in that week and, among all individuals who answered yes to this question, the number of hours worked. Among workers who reported working positive hours in the reference week, the PNAD reports whether individuals' primary economic activity was as an employee earning a wage or salary. Second, the PNAD provides information on the earnings of wage and salary workers: specifically, among all individuals who worked positive hours during the reference week, and who report that their primary activity during this week was as an employee in a job (the "primary" job), the PNAD reports the average yearly wage from that primary job. I use these variables to construct measures of hourly earnings among wage and salary earners. The PNAD also provides information on standard individual covariates, including sex, age and years of schooling. I construct a measure of (Mincerian) potential labor market experience similar to that used in Katz and Murphy:

$$\text{potential experience} = \min(\text{age} - \text{years of schooling} - 5, \text{age} - 17)$$

Particularly-important for the Brazilian context, this measure prevents people with low levels of education from having more-than-plausible years of experience (e.g., a 17 year old with 2 years of schooling is captured as having has 0 years of plausible experience, not 10).⁷⁴

While almost all work in the paper is drawn from the PNAD, in Figure 10 I use the 1940 and 1950 U.S. Household Censuses, as well as the 1991 and 2010 Brazilian Household Censuses in order to compare the "Great Compression" of the wage distribution in the U.S. to wage changes in Brazil over the last 20 years. Because these data sources are only used briefly, I document data processing

⁷⁴This measure differs from the original Katz and Murphy measure of potential experience, in which the quantity age - years of schooling - 7 is used, to account for the fact that primary schooling typically begins at age 5 in Brazil.

procedures in the notes to Figure 2.

II.B.2 The Price and Quantity Samples

The primary analysis of relative wages and supplies in this paper considers male workers between the ages of 17 and 52 during the 1995-2011 earnings years.^{75,76} For these years, I follow Katz and Murphy in constructing separate samples to measure wages and supplies - the "price" and "quantity" samples. As a first step, I divide male workers into 90 separate groups, defined by 5 narrow schooling groups by 18 narrow potential experience groups. The 5 narrow schooling groups consider men with 0, 1-4 (primary), 5-8 (lower secondary), 9-11 (upper secondary), and 12+ (college and above) years of schooling. The 18 potential experience groups divide men with between 0 and 35 years of potential experience into 18 separate 2-year bins.⁷⁷

In the price sample, intended to capture the wages of full-time (but not necessarily full-year)⁷⁸ workers, men are required to be an employee in their primary economic activity, to have worked at least 35 hours in that job in the reference week, to be between the ages of 17 and 52, and to have non-missing values for years of schooling and wages in the data. These relatively-restrictive assumptions are intended to reduce the influence of sample composition on changes in measures of wages across time. In the quantity sample, by contrast, workers are required only to have worked in the reference week (but not over a certain number of hours in the reference week, and not only as a wage and salary earner), to be between the ages of 17 and 52, and to have non-missing values for years of schooling in the data.

⁷⁵In Table 1 and Figure 1, I examine both male and female workers of a slightly-larger age range of workers (17-56), and I construct price and quantity samples similar to those defined in this section in order to generate measures of employment shares in efficiency units. However, all subsequent analyses of relative wages and supplies employ the price and quantity samples exactly as defined in this section.

⁷⁶While recent analyses of relative wages in the U.S., such as Autor et al. (2008), consider men as old as 64, here I restrict my attention to workers below the age of 53 since the Brazilian male workforce is considerably younger than the U.S. workforce. Across the years considered, the fraction of the male Brazilian workforce over the age of 52 is quite low.

⁷⁷By restricting the age of males to between 17 and 52, and using the measure of Mincerian potential experience described above, I implicitly consider workers with between 0 and 35 years of potential experience.

⁷⁸Across the 1995-2011 earnings years, the PNAD does not contain a measure of weeks worked per year, necessary to restrict the sample to full-year workers.

II.B.3 Measures of Mean Wages and Supplies of Demographic Groups

Following Katz and Murphy, I jointly use the price and quantity samples to construct (i) composition-adjusted measures of mean wages of broad demographic groups, and (ii) measures of the supplies of broad demographic groups, measured in efficiency units. First, for a given broad demographic group, the mean wage in a given year is a weighted average of the underlying narrow schooling-by-potential experience cells from the price sample, weighted by the associated average employment shares of these underlying groups, drawn from the quantity sample.⁷⁹ Second, the supply of a broad demographic group, as measured in efficiency units, represents the weighted sum of the underlying narrow schooling-by-potential experience cells from the quantity sample, where the weights are the associated average relative mean wages of these underlying groups, drawn from the price sample. For each schooling-by-potential experience cell, these average relative mean wages represent the efficiency units of labor supplied in every period by workers of the given group.

II.C *Patterns of Wage Inequality, Relative Wages and Supplies*

In this section, I examine patterns of wage inequality, relative wages and relative supplies across the 1976-2011 earnings years in Brazil. I begin by documenting the smooth, long-run expansions of schooling and labor market experience in the Brazilian workforce. I then examine measures of wage inequality across all earnings years from 1976 to 2011. I find that, independent of the chosen price deflator, real wage data before the end of hyperinflation in Brazil in 1994 contain implausibly-large short-run movements in both absolute and relative wages. As a result, in the rest of the paper, I examine wages and supplies of male workers exclusively in post-hyperinflation Brazil: from 1995 to 2011. I examine changes in real wages and wage inequality across this period, and I show that two distinct periods emerge. First, from 1995 to 2002, while per capita GDP growth stagnated and the real incomes of most men fell, all segments of the wage distribution compressed modestly, with the exception of the left tail, which compressed dramatically. Second, from 2002-2011, as per capita GDP growth grew by roughly 2.7% annually and real wages increased at all percentiles, the log wage distribution experienced compression that was monotonically-decreasing and nearly-linear in wage

⁷⁹These mean wage measures differ in at least two ways from more-recent work on relative wages in the U.S. First, rather than using mean wages, studies such as Autor et al. (2008) examine mean log wages. Second, rather than using simple means, these studies control for composition effects using a regression methodology. As noted earlier, here I follow existing studies on Brazilian and Latin American relative wages by more-closely following the work of Katz and Murphy (1992).

percentile. Finally, simple models of schooling wage differentials, originally-developed for the U.S. labor market and recently-applied to the labor markets of Brazil and other Latin American countries suggest that changes in the relative wages of groups of workers defined solely by level of schooling may be an important correlate of changes in raw wage inequality: in the final section, I document changes in relative wages and supplies between these groups.

TABLE 9
Employment Shares Among All Workers, by Demographic Group: 1977-2009

	1977	1981	1985	1989	1993	1997	2001	2005	2009
A. Efficiency Units									
Gender:									
Men	0.80	0.78	0.76	0.75	0.73	0.71	0.70	0.68	0.67
Women	0.20	0.22	0.24	0.25	0.27	0.29	0.30	0.32	0.33
Yrs. Schooling:									
0	0.12	0.09	0.08	0.06	0.06	0.04	0.03	0.02	0.02
1-4	0.35	0.31	0.28	0.24	0.21	0.17	0.13	0.11	0.07
5-8	0.19	0.19	0.20	0.20	0.21	0.21	0.20	0.18	0.15
9-11	0.14	0.17	0.19	0.22	0.23	0.27	0.32	0.36	0.38
12+	0.19	0.23	0.25	0.28	0.28	0.30	0.32	0.33	0.38
Yrs. Potential Experience:									
0-7	0.23	0.21	0.20	0.18	0.16	0.15	0.15	0.15	0.14
8-15	0.27	0.28	0.29	0.28	0.28	0.25	0.23	0.23	0.24
16-23	0.22	0.23	0.24	0.27	0.28	0.29	0.28	0.26	0.25
24-31	0.18	0.17	0.17	0.17	0.19	0.22	0.23	0.23	0.24
32-39	0.11	0.10	0.10	0.10	0.09	0.10	0.11	0.13	0.14
B. Total Hours Worked									
Gender:									
Men	0.73	0.72	0.70	0.68	0.66	0.65	0.63	0.62	0.61
Women	0.27	0.28	0.30	0.32	0.34	0.35	0.37	0.38	0.39
Yrs. Education:									
0	0.23	0.19	0.17	0.15	0.14	0.11	0.09	0.07	0.06
1-4	0.45	0.43	0.40	0.37	0.34	0.29	0.24	0.21	0.15
5-8	0.19	0.20	0.22	0.24	0.26	0.28	0.27	0.26	0.23
9-11	0.08	0.12	0.14	0.16	0.17	0.21	0.28	0.33	0.38
12+	0.05	0.06	0.07	0.08	0.09	0.10	0.12	0.14	0.17
Yrs. Potential Experience:									
0-7	0.33	0.31	0.30	0.28	0.26	0.24	0.23	0.23	0.21
8-15	0.24	0.25	0.26	0.25	0.26	0.25	0.24	0.24	0.24
16-23	0.19	0.19	0.20	0.22	0.22	0.23	0.23	0.23	0.22
24-31	0.14	0.15	0.14	0.15	0.16	0.17	0.18	0.19	0.19
32-39	0.10	0.10	0.09	0.10	0.10	0.10	0.11	0.12	0.13

TABLE 10
Log Wage Inequality Among All Workers: 1977-2009

	1977	1981	1985	1989	1993	1997	2001	2005	2009
	A. Log Hourly Wage								
S.D.	0.93	0.94	0.96	1.02	0.97	0.90	0.86	0.79	0.73
Percentile differential:									
90-10	2.21	2.38	2.36	2.48	2.38	2.20	2.04	1.85	1.63
75-25	1.05	1.14	1.17	1.33	1.15	1.19	1.01	0.87	0.79
90-50	1.32	1.35	1.40	1.47	1.35	1.28	1.23	1.17	1.07
50-10	0.89	1.03	0.96	1.02	1.03	0.91	0.80	0.68	0.55
75-50	0.62	0.63	0.68	0.72	0.66	0.61	0.54	0.51	0.48
50-25	0.42	0.51	0.50	0.61	0.49	0.59	0.47	0.36	0.30
	B. Log Monthly Wage								
S.D.	0.89	0.91	0.92	0.98	0.94	0.87	0.83	0.77	0.71
Percentile differential:									
90-10	2.11	2.30	2.23	2.33	2.30	2.10	1.85	1.65	1.46
75-25	1.01	1.10	1.20	1.34	1.10	1.14	1.05	0.91	0.83
90-50	1.26	1.27	1.32	1.38	1.32	1.23	1.19	1.12	1.07
50-10	0.85	1.03	0.91	0.94	0.98	0.87	0.66	0.52	0.39
75-50	0.61	0.59	0.67	0.68	0.63	0.56	0.54	0.50	0.48
50-25	0.39	0.51	0.53	0.66	0.47	0.58	0.51	0.41	0.35

Figure 9: Employment Shares of Education and Potential Experience Groups Among All Workers, in Efficiency Units: 1976-2011

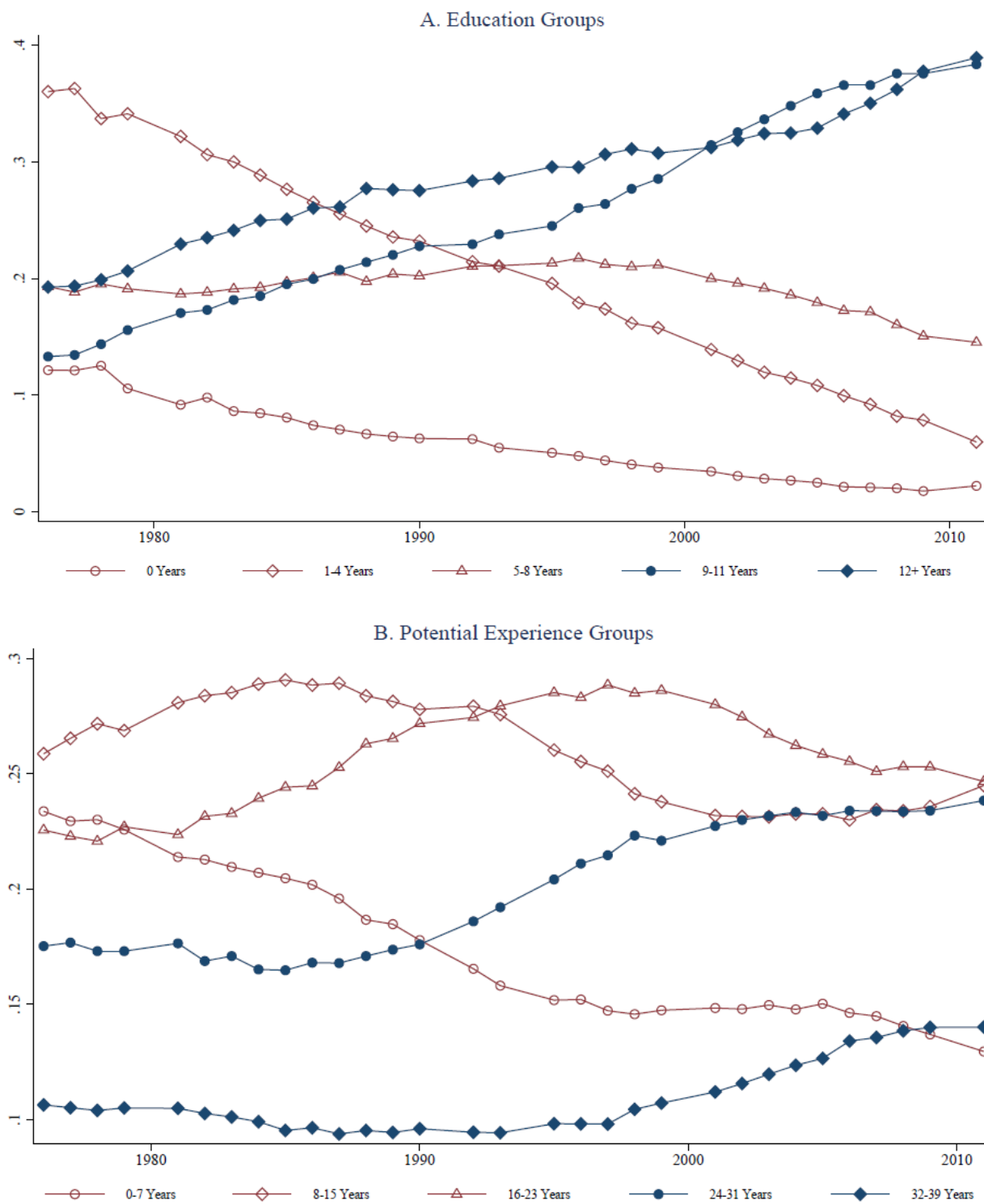
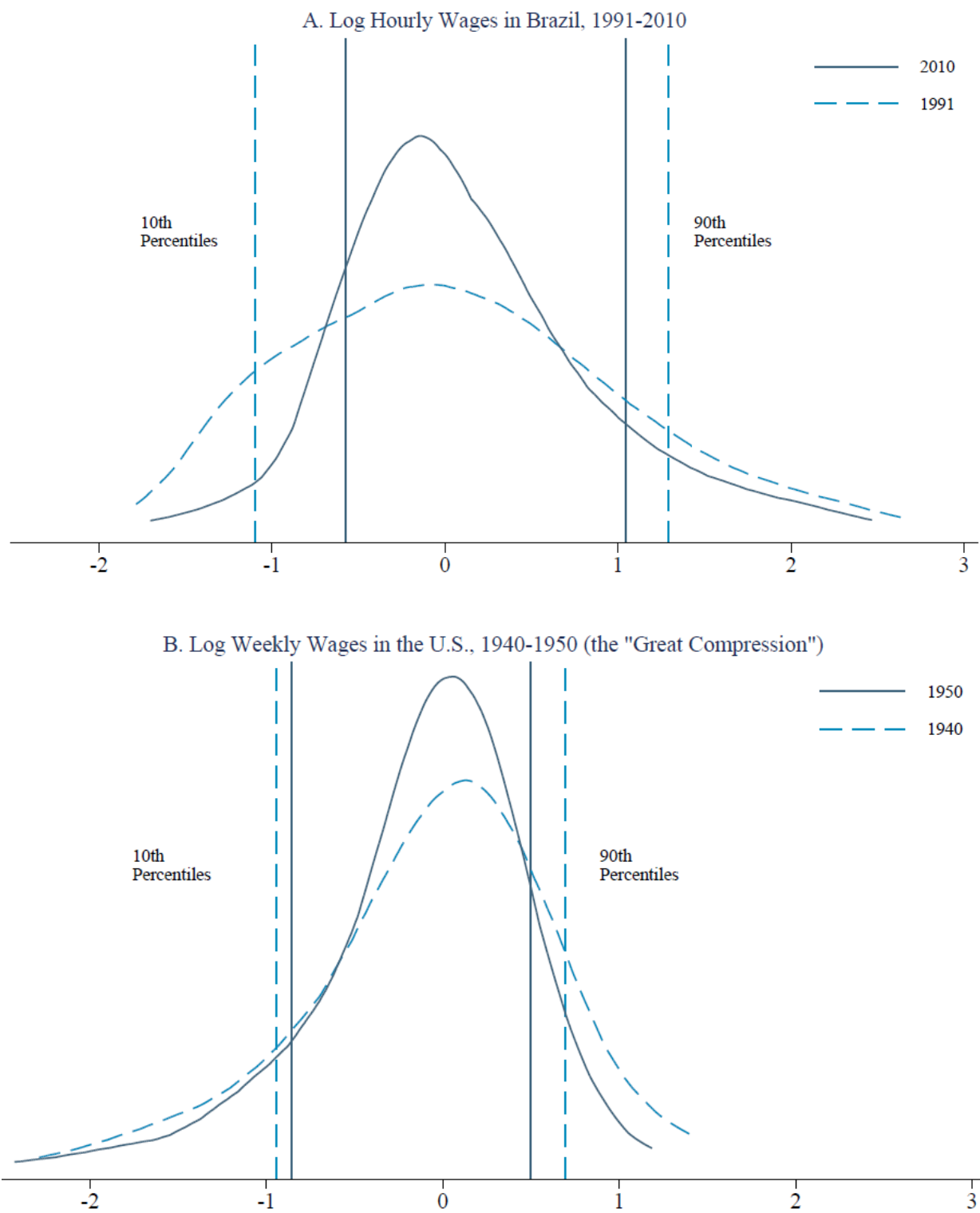


Figure 10: Log Wage Densities (Centered on their Medians) for Male Workers

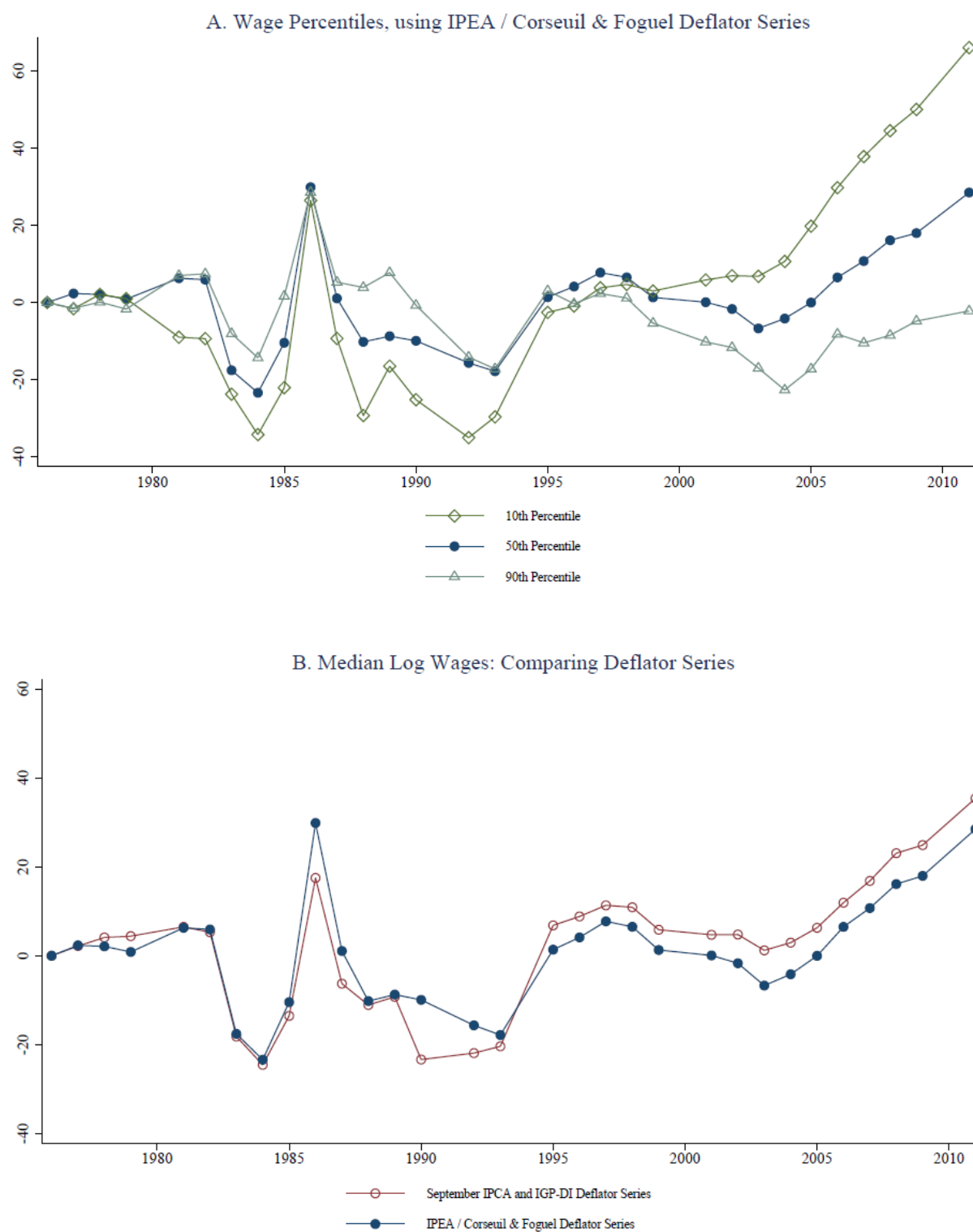


II.C.1 The Smooth Expansion of Schooling and Experience: 1976-2011

In his study of educational expansion in Brazil, Plank (1987) finds a clear long-run trend break in school enrollment rates in nearly every state of Brazil in the 1950s; separately, Friedrich de Amaral (2007) documents a striking decline in historical fertility rates: from roughly 6 in the 1950s to roughly 2.5 in the 1990s. Consistent with these basic facts, Table 9 and Figure 9 document smooth, long-run increases in the level of schooling (potentially associated with long-run increases in enrollment rates) and labor market experience (potentially associated with long-run changes in fertility) among employed workers in Brazil from 1976-2011.

Inspection of Panel B of Table 9 reveals a dramatic expansion of the level of schooling, and a moderate expansion in the level of potential experience among employed workers. For example, the total hours supplied by workers with some upper-secondary education (at least 9 years of education) represented only 13% of total hours supplied in 1977, but quickly grew to 55% in 2009. The percentage of total hours supplied by workers with at least 24 years of potential labor market experience grew relatively-moderately: from 24% in 1977 to 33% in 2009. Both panels of Table 1, as well as the employment shares, in efficiency units, plotted in Figure 1, document the smoothness of these changes. Panel A of Figure 1, for example, documents the nearly-linear and rapid increases of the shares of employment of workers with upper-secondary education (9-11 years of schooling) and tertiary education (12+ years), as well as a nearly-linear and equally-rapid decline in the share of employment of workers with only primary education (1-4 years). Panel B of Figure 9 similarly reveals smooth movements both towards labor with at least 24 years of potential labor market experience, and away from labor with fewer than 8 years of experience.

Figure 11: Cumulative Change in Measures of Log Wage Percentiles Among All Workers (times 100): 1976-2011



II.C.2 Real Wages, 1976-2011: Are Measured Changes Before 1995 Plausible?

While movements in levels of schooling and potential experience are smooth across the entire 1976-2011 period, movements in measures of real and relative wages are not. A large rise, from 1976-1989, then fall, from 1989-2011, in measures of log real wage inequality drawn from reported wages in the PNAD (the only yearly source of wage data in Brazil that spans this entire period), has been extensively-documented and has generated substantial interest in the literature (Barros et al., 2009; Ferreira et al., 2006; Lustig et al., 2011; 2012; Sotomayor, 2004). Moreover, applications of standard supply and demand frameworks in Brazil have examined changes in relative wages that include years before the end of hyperinflation in 1994 (Blom and Velez, 2004; Gasparini et al., 2011). However, potentially due to dramatic changes in inflation between 1980 and 1994, short-run movements in real and relative wages before 1995 appear to be implausible.

First, Table 10 reviews the measured rise and fall in log wage inequality across the 1976-2011 earnings years. Every measure the variance of both the log hourly and log monthly real wage distributions documented in the PNAD increased from 1977-1989, then fell dramatically from 1989-2009. For example, the p90-10 gap of the log monthly wage distribution increased by 0.22 across the 1977-1989 earnings years, then decreased by 0.89 from 1989 to 2009. In the next section, I will argue that much of the latter measured decrease is plausibly accurately-measured, and is composed in large part by changes that occurred after the end of hyperinflation in 1994. Figure 10 presents changes in the log hourly real wage distribution across the 1991-2010 earnings years among male workers in Brazil (using household census data), compared to changes in the distribution of log real weekly wages among male workers during the "Great Compression" in the U.S. (Goldin and Margo, 1992). At least two facts are striking: first, the overall variance of the Brazilian log wage distribution is significantly larger than in the U.S. (the measured p90-10 gap is 2.39 in Brazil in 1991, compared to 1.63 in the U.S. in 1940). Second, compression of the distribution of log hourly wages in Brazil comes primarily from its lower tail, in contrast to changes in the U.S.⁸⁰ I return to this second fact in the next section.

While Figure 10 primarily reflects measured wage changes after the end of hyperinflation in 1994, inspection of Figure 11 reveals the significant difficulty involved in using wage observations between

⁸⁰Likely due to narrower age criteria and different trimming assumptions, the log wage densities estimated in Panel B of Figure 2 are slightly different from those measured in Goldin and Margo (1992), who estimate that during the 1940-1950 period, among white male wage and salary earners between the ages of 19 and 64, the p90-10 gap decreased from 1.414 to 1.06, with the p90-50 gap decreasing by 0.22 and the p50-10 gap decreasing by 0.15.

the years of 1980 and 1994. First, Panel B of Figure 11 compares changes in median log wages. In this panel, I compare medians computed using a deflator series especially-designed for use with the PNAD (Corseuil and Foguel, 2002; updated by the Instituto de Pesquisa Econômica Aplicada) to medians computed using other standard deflator series (the IPCA and IGP-DI deflators). In either case, the measured level of log real wages undergoes implausible short-run jumps. As computed by the Corseuil and Foguel deflator, for example, measured median wages changed by roughly -23% from 1982-1993, 40% from 1985-1986, -29% from 1986-1987, and 19% from 1993-1995.

In principle, if measures of all points on the log wage distribution were identically-affected by measurement error before 1994, *relative* wages could still be analyzed during across this period. However, Panel A of Figure 11 shows that movements in relative wages also exhibit implausibly-large jumps. As measured in the PNAD, the p50-10 gap increased by 0.15 between 1979 and 1981, and the p90-10 gap compressed by roughly 0.22 between 1985-1986, only to increase by 0.31 the following year.

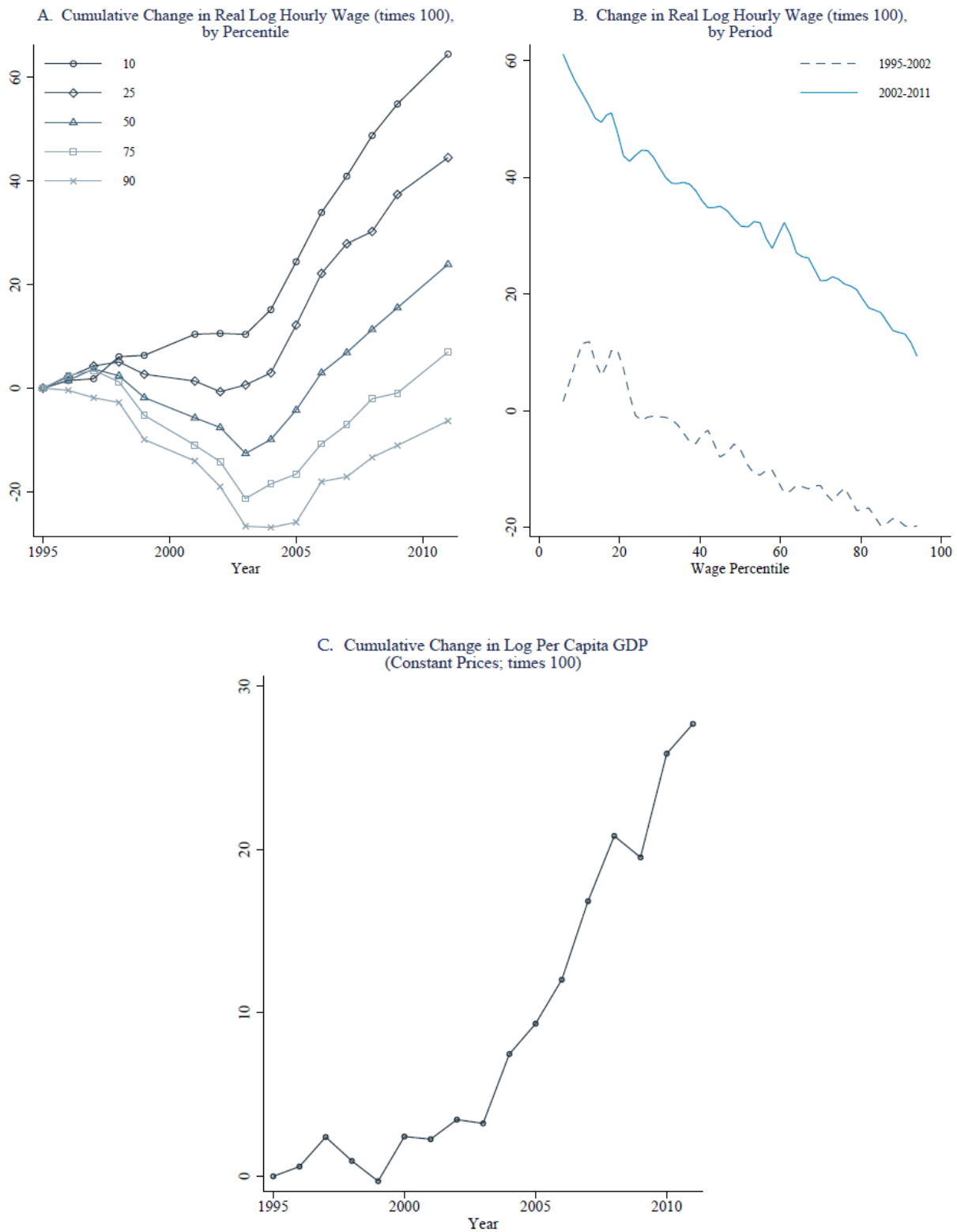
TABLE 11
Log Hourly Wage Inequality Among Male Workers: 1995-2011

A. Log Wage Inequality Measures: Level in Each Year						
	1995	1998	2002	2005	2008	2011
S.D.	0.91	0.89	0.85	0.78	0.74	0.70
Percentile differential:						
90-10	2.28	2.22	2.00	1.83	1.68	1.59
75-25	1.19	1.14	1.07	0.89	0.88	0.82
90-50	1.30	1.29	1.19	1.11	1.09	1.02
50-10	0.98	0.93	0.81	0.72	0.59	0.57
75-50	0.61	0.60	0.56	0.48	0.51	0.45
50-25	0.58	0.54	0.51	0.41	0.37	0.38
B. Log Wage Inequality Measures: Annual Change						
	1995-2002	2002-2011	1995-2011			
S.D.: Annual % Change	-0.93%	-1.98%	-1.45%			
Percentile differential:						
Annual Change (times 100)						
90-10	-3.9	-4.6	-4.3			
75-25	-1.6	-2.7	-2.3			
90-50	-1.5	-1.9	-1.7			
50-10	-2.4	-2.7	-2.6			
75-50	-0.7	-1.3	-1.0			
50-25	-0.9	-1.5	-1.2			

TABLE 12
Changes in Real Hourly Wages Among Male Workers, 1995-2011

Group	A. Change in Log Mean Real Hourly Wage (multiplied by 100)			B. Change in Log Mean Real Hourly Wage: "Composition- Adjusted" Component (multiplied by 100)			C. Change in Log Mean Real Hourly Wage: "Composition Effect" Component (multiplied by 100)		
	1995- 2002	2002- 2011	1995- 2011	1995- 2002	2002- 2011	1995- 2011	1995- 2002	2002- 2011	1995- 2011
All Men	-11.1	21.2	10.1	-24.1	6.9	-17.2	13.0	14.3	27.3
<u>Yrs. Schooling</u>									
Narrow Groups:									
0	-2.1	55.1	53.0	-3.2	54.0	50.8	1.0	1.1	2.2
1-4	-19.7	29.3	9.7	-23.1	23.7	0.6	3.4	5.7	9.1
5-8	-22.9	19.8	-3.1	-29.3	16.1	-13.2	6.4	3.7	10.1
9-11	-30.8	4.9	-25.9	-29.3	0.8	-28.4	-1.5	4.1	2.5
12+	-20.2	-13.8	-34.0	-17.2	-12.3	-29.5	-3.0	-1.5	-4.6
Broad Groups:									
0-8 "Non-Upper-Secondary"	-17.2	27.5	10.4	-24.7	23.2	-1.5	7.6	4.3	11.9
9+ "Upper-Secondary"	-30.8	-0.5	-31.3	-23.6	-5.3	-28.9	-7.2	4.8	-2.4
<u>Yrs. Potential Experience</u>									
0-7 "Inexperienced"	-6.4	33.3	26.9	-24.0	24.5	0.5	17.5	8.8	26.4
8-17	-15.4	23.5	8.1	-23.3	7.6	-15.7	7.9	15.9	23.8
18-35 "Experienced"	-11.7	11.6	-0.1	-24.5	1.5	-23.0	12.8	10.2	22.9
<u>Education-by-Potential Experience:</u>									
0-8 Yrs. Schooling:									
0-7 Yrs. Exp.	-15.2	39.7	24.4	-20.1	37.9	17.8	4.8	1.8	6.6
18-35 Yrs. Exp.	-22.2	21.1	-1.1	-27.5	19.3	-8.3	5.4	1.8	7.1
9+ Yrs. Schooling:									
0-7 Yrs. Exp.	-28.3	18.2	-10.1	-26.6	14.1	-12.5	-1.7	4.1	2.4
18-35 Yrs. Exp.	-27.3	-13.7	-40.9	-22.4	-12.6	-35.0	-4.9	-1.0	-5.9

Figure 12: Changes in Real Hourly Wages Among Males and Per Capita GDP: 1995-2011



II.C.3 Changes in Real Wages and Wage Inequality Among Males, 1995-2011

To ensure that my analysis is free of such obvious measurement error, in the rest of the paper I examine only data from post-hyperinflation Brazil: 1995-2011. Further, because my analysis depends on proper measurement of both schooling and labor market experience, I examine only male workers.⁸¹ Table 11 and Figure 12 develop the basic facts of log hourly wage inequality among Brazilian male wage and salary earners in Brazil from 1995 to 2011.

1995-2002: Stagnant GDP Growth; Real Wage Losses; Left Tail Compression From 1995-1998, most men saw their real wages increase. Over this period, inequality decreased only slightly: Panel A of Figure 12 shows that all but the most-skilled/lucky men (those men at the top of the wage distribution) saw their wages rise modestly, while those at the top saw slight real wage declines; Panel A of Table 11 shows that, overall, most measures of wage inequality remained relatively-unchanged from 1995-1998. These modest wage movements coincided with a similarly-modest 0.3% annual rate of yearly per-capita GDP growth over the three-year period.

From 1998-2002 however, Panel A of Figure 12 reveals that a compression of nearly all segments of the wage distribution coincided with a drop in real wages for all but the least-skilled/lucky men (those men at the bottom of the wage distribution). Note that real wages begin to recover in 2002 for the 25th percentile, in 2003 for the 50th and 75th, and in 2004 for the 90th percentile. Inspection of Panel B of Figure 12 makes clear that the compression of the wage distribution from 1995-2002 is most-pronounced among the less-skilled/lucky: only men beneath the 20th percentile of the wage distribution made positive real wage gains over this period. This left tail compression is reflected in Panel B of Table 3: while the middle of both the top and the bottom of the wage distribution compressed relatively-evenly across this period (the p75-50 and p50-20 gaps compressed by 0.7 and 0.9 log points annually, respectively), figures in Table 11 imply that the p90-75 gap compressed by only 0.8 log points annually, while the p25-10 gap compressed by nearly twice as much: 1.5 log points annually over this period.

⁸¹Mincerian potential labor market experience may not accurately-capture the level of actual labor market experience among women if female attachment to the labor force is weak; moreover, if female labor force participation rates increase over time, changes in potential labor market experience may mis-measure actual changes in labor market experience across entry cohort. In focusing exclusively on the supply and demand for male labor, I follow a long literature on wage inequality in the U.S. (Card and Lemieux, 2001; Goldin and Margo, 1992; Juhn et al., 1993; Murphy and Welch, 1992).

2002-2011: Robust GDP Growth; Real Wage Gains; Linear Compression In contrast to the relative stagnation of the 1995-2002 period, from 2002 to 2011, real wages for all groups made substantial gains, and inequality decreased further as these decreases were monotonically-, and nearly-linearly decreasing in the percentile of the wage distribution: Panel B of Figure 12 shows that, over this period, real wages increased by approximately 15% for the 95th percentile of the wage distribution, while wages increased by approximately 60% for the 5th percentile of the wage distribution. Panel B of Table 12 shows that the gap between the 90th and 10th percentiles of the wage distribution closed by roughly 4.6% annually: roughly 1.1% per year faster than the analogous gap for weekly log wages among white males during the Great Compression of the U.S.⁸²

Taken together, these figures point to two distinct periods of changes in real wage inequality. First, from 1995-2002, most men experienced real wage losses as output in the economy stagnated. However, while wage decreases were relatively-similar for all but the men at the lowest part of the wage distribution, real wage gains of the least-skilled/lucky workers (those below the 20th percentile of the wage distribution) increased substantially. Second, 2002-2011, there is a near-linear compression of the log wage distribution as per-capita GDP increased by roughly 25%.

II.C.4 Changes in Relative Wages and Supplies of Schooling Groups, 1995-2011

In the U.S. labor market, changes in raw wage inequality are closely-related to changes in schooling wage differentials: roughly 60% of the post-1980 increase in U.S. wage inequality can be accounted for by increases in the returns to post-secondary education (Firpo et al., 2007; Goldin and Katz, 2007; Lemieux, 2006a). A number of recent papers have directly-applied models of schooling wage differentials to the Brazilian labor market, as well as other labor markets in Latin America (Cruces et al., 2012; Galiani, 2009; Gasparini et al., 2011; Manacorda et al., 2010). In this section, I document changes in the relative wages and supplies of groups of workers defined by their level of schooling. In the next section of the paper, I then examine whether a simple model of schooling premia is appropriate for the Brazilian context.

⁸²Goldin and Margo (1992) estimate that during the 1940-1950 period, among white male wage and salary earners between the ages of 19 and 64, the p90-10 gap decreased from 1.414 to 1.06: a total decrease of 0.354 over the 10-year period.

Relative Wages Table 12 presents changes in measures of real hourly wages for male workers. Wage measures of interest are, in keeping with the original analysis of Katz and Murphy, (i) adjusted for changes in the composition of underlying demographic groups, and (ii) expressed as mean wages, rather than log wages.⁸³ Panels B and C reflect the decomposition, detailed in Appendix A.6, of raw change in the log mean real hourly wage (Panel A) into two components: a "composition-adjusted" component (Panel B) and a "composition effect" component (Panel C). Note that, due to (i) large schooling wage differentials in Brazil, combined with (ii) rapid movement of workers from lower to higher levels of education during this period (both documented below), shifts of workers across schooling groups account for large increases in the raw changes of Panel A. For example, consistent with movements in median wages presented in Panel A of Figure 4, Panel A of Table 12 shows that raw mean real hourly wages increased by roughly 21.2% from 2002-2011; Panels B and C show that roughly 14.3% is "accounted for" by changes in the composition of the workforce, while only 6.9% is accounted for by "composition-adjusted" changes. In the rest of this paper, I consider only composition-adjusted wage changes, such as those displayed in Panel B of Table 4.

Inspection of Panel B of Table 12 shows that changes in composition-adjusted wages of schooling groups can be broken down into two distinct periods. First, from 1995-2002, schooling wage differentials were relatively stable. Consider first two broad and comprehensive schooling groups: workers without any upper-secondary education (fewer than 8 years of schooling) and workers with at least some upper-secondary education (at least 9 years of schooling). From 1995-2002, Panel B shows that these two groups both saw real wage losses of approximately 24%. Broken down into "narrow" schooling groups, this pattern becomes slightly more-complicated: workers with primary (1-4 years of schooling), lower-secondary (5-8 years of schooling) and upper-secondary (9-11 years of schooling) all experienced real wage losses of roughly 23-29%, while groups on the extreme ends of the schooling distribution (either no years of schooling or tertiary education) increased more-quickly.

From 2002-2011 however, composition-adjusted schooling wage differentials - across both the broad and narrow schooling groups displayed in Table 12 - compressed dramatically as less-educated workers saw composition-adjusted real wage gains. First, while real wages of the broad group of non-upper-secondary men increased by roughly 23% from 2002-2011, real wages for upper-secondary-educated

⁸³ As mentioned in the section on data and empirical measures above, while more-recent literature on changes in relative wages in the U.S. examines log wages, rather than mean wages, of demographic groups, in this paper I follow the literature on Brazilian and Latin American relative wages, which follows the original empirical choices of Katz and Murphy.

workers decreased, on average, by roughly 5%. Second, composition-adjusted real wage changes among the narrow schooling groups in Table 12 are monotonically-decreasing with formal schooling over this period, with all but the most-educated workers seeing real wage gains: these wage changes are on the order of those (raw changes) displayed for percentiles of the overall log wage distribution in Panels A and B of Figure 4: while Panel B of Figure 12 shows that workers at the 90th and 10th wage percentiles saw their real wages rise by roughly 15% and 60%, respectively, between 2002 and 2011, Panel B of Table 12 shows that workers with 12+ and 0 years of schooling saw their composition-adjusted real wages change by roughly -12% and 54%, respectively.

TABLE 13
Employment Shares Among Male Workers, in Efficiency Units: 1995-2011

Group	A. Employment Share			B. Average Annual Change in Log Employment Share (times 100)		
	1995	2002	2011	1995- 2002	2002- 2011	1995- 2011
<u>Yrs. Schooling</u>						
Narrow Groups:						
0	0.07	0.04	0.03	-6.2	-2.6	-4.2
1-4	0.23	0.16	0.08	-5.3	-8.1	-6.9
5-8	0.24	0.23	0.18	-0.6	-2.8	-1.9
9-11	0.22	0.31	0.39	4.8	2.4	3.5
12+	0.24	0.26	0.32	0.8	2.6	1.8
Broad Groups:						
0-8 "Non-Upper-Secondary"	0.54	0.43	0.29	-3.1	-4.5	-3.9
9+ "Upper-Secondary"	0.46	0.57	0.71	2.9	2.5	2.7
<u>Yrs. Potential Experience</u>						
0-7 "Inexperienced"	0.15	0.15	0.13	-0.5	-1.9	-1.3
8-17	0.33	0.30	0.32	-1.3	0.8	-0.1
18-35 "Experienced"	0.52	0.55	0.55	0.9	0.0	0.4
<u>Education-by-Potential Experience:</u>						
0-8 Yrs. Schooling:						
0-7 Yrs. Exp.	0.10	0.06	0.03	-6.2	-7.8	-7.1
18-35 Yrs. Exp.	0.27	0.24	0.18	-1.8	-2.9	-2.4
9+ Yrs. Schooling:						
0-7 Yrs. Exp.	0.06	0.09	0.09	6.2	1.1	3.3
18-35 Yrs. Exp.	0.24	0.31	0.37	3.4	1.9	2.5

Relative Supplies In contrast to schooling wage differentials, which were relatively stable from 1995-2002, then compressed dramatically from 2002-2011, Table 13 documents smooth increases in the relative supply of more-educated workers across the entire 1995-2011 period. For example, the employment share, in efficiency units, of workers with upper-secondary education increased from 0.46 in 1995 to 0.57 in 2002 to 0.71 in 2011; figures from Panel B of Table 13 imply that the associated annual change in the log relative supply of workers with upper-secondary, relative to non-upper-secondary education increased by roughly 5.9% annually across the 1995-2002 period and 7% annually across the 2002-2011 period.

The patterns of relative wages and supplies of schooling groups in Brazil from 1995-2011 captured in Tables 4 and 5 are suggestive: while relative supplies of more-educated workers increased dramatically and smoothly across the entire 1995-2011 period, schooling differentials across all segments of the schooling distribution experienced greater narrowing during the 2002-2011 period than during the 1995-2002 period. One plausible hypothesis is that, over the 2002-2011 earnings years, growth in the relative demand for more-educated workers in Brazil slowed. In the next section, I examine whether observed time series of relative wages and relative supplies are consistent with a simple version of this hypothesis.

II.D Modeling Relative Wages in Brazil, 1995-2011

Changes in the relative wages and supplies of more-educated workers uncovered in the previous section appear to suggest a simple supply and demand story: smooth increases in the relative supply of more-educated workers, in combination with a sharp slowdown in relative demand in 2002, may generate the large decline in schooling wage differentials observed across the 2002-2011 period. In this section, I examine whether a standard formalization of this supply and demand story is consistent with the data.

I begin by constructing a simple model of the schooling wage premium, which I measure as the ratio of the composition-adjusted mean wage of workers with some upper-secondary education to workers without such education. This model is intended to match the theoretical and empirical choices of standard models of schooling premia originally-developed for the U.S. labor market (Card and Lemieux, 2001; Goldin and Katz, 2008; 2009; Katz and Murphy, 1992) and recently-used to analyze the decline of schooling premia in Brazil and other Latin American countries in the 2000s (Cruces et al., 2012; Galiani, 2009; Gasparini et al., 2011; Manacorda et al., 2010). Within this standard model, I show

that if demand shifts are constant within schooling groups, (i) the time series of schooling premia should be similar for all potential experience groups, and (ii) potential experience wage differentials should be approximately-constant within a schooling group. I show that both of these predictions are inconsistent with observed changes in wage differentials, and cannot be easily-explained by changes in relative supplies.

As an alternative to direct-application of standard U.S. frameworks to the Brazilian context, I sketch a simple alternative approach. I propose a model that groups workers not only by their level of schooling, but also by their level of labor market experience. I compare three types of workers: low-skill (less-educated, less-experienced), high-skill (more-educated, more-experienced), and middle-skill (either more-educated or more-experienced, but not both). I flexibly-apply this alternative framework to examine the determinants of changes over time in relative wages between these groups in Brazil from 1995-2011.

II.D.1 Are Data Consistent with a Standard Model of the Schooling Premium?

In this section, I examine whether observed changes in relative wages and supplies are consistent with a standard supply and demand model of the schooling premium. Although authors who have investigated schooling premia in Brazil and other Latin American countries have examined a variety of different schooling wage differentials and have employed slightly different variations of standard frameworks (Card and Lemieux, 2001; Katz and Murphy, 1992), nearly all such analyses have assumed that workers within a given schooling group face identical demand shifts (Blom and Velez, 2004; Cruces et al., 2012; Gasparini et al., 2011; Manacorda et al., 2010; Montes Rojas, 2006). Here I take as my measure of the schooling premium the ratio of the composition-adjusted mean wage of workers with some upper-secondary education to workers without such education,⁸⁴ and I construct a simple standard supply and demand model in which this assumption, and other standard assumptions, are retained. I show that this model has two key predictions, and I test these predictions empirically.

⁸⁴This choice contrasts with the standard schooling wage differential studied in the U.S., the college wage premium. I make this distinction for two reasons. First, in Brazil only a very small fraction of the labor force completes any tertiary education before entering the labor market (e.g., only 10% in 1997, according to Panel B of Table 1), suggesting that upper-secondary schooling, simply by virtue of the share of the population that obtains this level of schooling, may be more-relevant. Second, unlike the U.S., in which a large literature argues that the skills gained by attending college are of first-order importance to the labor market (Becker, 1964; Goldin and Katz, 2008; Mincer, 1974), in Brazil relatively-little literature points towards a specific schooling category as particularly-relevant in the labor market.

A Standard Model: Two Key Predictions Consider a simple model in which aggregate production is assumed to be undertaken using two inputs: labor supplied by workers who completed some upper-secondary education before labor market entry (henceforth "upper-secondary workers"), and workers who did not complete any upper-secondary education before labor market entry (henceforth "non-upper-secondary workers"). Further assume, as is standard (e.g., Card and Lemieux, 2001), that (i) sub-groups of workers within each schooling group defined by their level of potential experience are associated with a constant elasticity of substitution, (ii) all workers within a schooling group face identical demand shifts, and (iii) aggregate production is approximated by a CES production function, so that aggregate production can be written as:

$$(46) \quad Y_t = \left[\alpha_{Nt} (A_{Nt} N_t)^{\frac{\sigma_s - 1}{\sigma_s}} + \alpha_{Ut} (A_{Ut} U_t)^{\frac{\sigma_s - 1}{\sigma_s}} \right]^{\frac{\sigma_s}{\sigma_s - 1}}$$

where Y_t is the quantity of aggregate output at time t , N_t is the quantity of non-upper-secondary labor at time t , U_t is the quantity of upper-secondary labor at time t , σ_s is the constant elasticity of substitution between these groups, and where for each schooling group $k \in \{N, U\}$, α_{kt} and A_{kt} are technology parameters of the share and labor-augmenting form, respectively. Assume that either schooling aggregate $k_t \in \{N_t, U_t\}$ can be written as:

$$(47) \quad k_t = \left[\sum_e (\pi_{ke} k_{et})^{\frac{\sigma_e - 1}{\sigma_e}} \right]^{\frac{\sigma_e}{\sigma_e - 1}}$$

where π_{ke} reflects the efficiency units of labor supplied in each period by workers of schooling group k and potential experience group e , k_{et} is the quantity of labor supplied by workers of schooling group k and potential experience group e at time t , and σ_e is the elasticity of substitution between potential experience groups. If labor markets are competitive so that workers are paid their marginal product, then for each potential experience group e :

$$\begin{aligned}
(48) \quad \ln \left(\frac{w_{Uet}}{w_{Net}} \right) &= \frac{\sigma_e - 1}{\sigma_e} \ln \left(\frac{\pi_{Ue}}{\pi_{Ne}} \right) \\
&+ \ln \left(\frac{\alpha_{Ut}}{\alpha_{Nt}} \right) + \frac{\sigma_s - 1}{\sigma_s} \ln \left(\frac{A_{Ut}}{A_{Nt}} \right) - \left[\frac{1}{\sigma_s} - \frac{1}{\sigma_e} \right] \ln \left(\frac{U_t}{N_t} \right) \\
&- \frac{1}{\sigma_e} \ln \left(\frac{U_{et}}{N_{et}} \right)
\end{aligned}$$

and for each schooling group k :

$$(49) \quad \ln \left(\frac{w_{ket}}{w_{ke't}} \right) = \frac{\sigma_e - 1}{\sigma_e} \ln \left(\frac{\pi_{ke}}{\pi_{ke'}} \right) - \frac{1}{\sigma_e} \ln \left(\frac{k_{et}}{k_{e't}} \right)$$

where w_{ket} is the wage of a worker of schooling group k and potential experience group e at time t . If all potential experience groups within a schooling group k are perfect substitutes ($\sigma_e \rightarrow \infty$) equations (48) and (49) immediately yield two key predictions. First, for a given potential experience group e :

$$(50) \quad \ln \left(\frac{w_{Uet}}{w_{Net}} \right) = \gamma_e + \gamma_t$$

Second, for a given schooling group k :

$$(51) \quad \ln \left(\frac{w_{ket}}{w_{ke't}} \right) = \gamma_k$$

That is, under the standard assumptions of equations (46) and (47) and perfect substitution within schooling groups, equation (50) implies that: (i) after adjusting for a constant, the time series of the log schooling premium should be identical for every potential experience group e , and (ii) the time series of any potential experience wage differential defined by the two groups e and e' should be fixed across time.

Figure 13: Schooling and Potential Experience Wage Differentials Among Male Workers, 1995-2011



Examining the Two Key Predictions Inspection of Panels B and C of Figure 13 shows that both of these predictions appear to be inconsistent with the data. First, consider Panel B of Figure 5, which compares change over time in the log schooling premium, defined as the log of the ratio of the composition-adjusted mean wage of workers with some upper-secondary education (9+ years) to workers without such education (0-8 years), for both inexperienced (0-7 years of potential experience) and experienced (18-35 years of potential experience) men. Equation (50) suggests that change over time in the log schooling premium should be identical for every potential experience group. However, Panel B (as well as figures from Table 4) shows that the log upper-secondary / non-upper-secondary wage premium takes two very different paths for inexperienced and experienced men. For inexperienced men, decrease over time in the schooling premium became moderately more-rapid in 2002: Panel B of Figure 5, as well as numbers from Table 4, show that across the 1995-2002 and 2002-2011 periods, the log schooling premium among inexperienced men decreased by roughly 7% and 24% over the period, or roughly 0.9% and 2.6% annually, respectively. By contrast, the log schooling premium among experienced men actually increased by roughly 5.2% (0.7% annually) from 1995-2002, then decreased by almost 32% (3.5% annually) from 2002-2011: a much-stronger trend-break.

Second, consider Panel C of Figure 5, which compares change over time in the log experience premium, defined as the log of the ratio of the composition-adjusted mean wage of experienced men (18-35 years of potential experience) to inexperienced men (0-7 years of potential experience), for both workers with some upper-secondary education (9+ years) to workers without such education (0-8 years). Equation (51) predicts that these wage differentials should be constant within both schooling groups. In contrast, Panel C of Figure 13 shows that the log experience premium among both upper-secondary-educated and non-upper-secondary educated workers was not constant, and in fact decreased by roughly 26% and 23% among workers with 0-8 and at 9+ years of schooling, respectively, between 1995 and 2011.

Can Movements in Supply Help Explain the Discrepancies? The two predictions of equations (50) and (51) above are based on the standard assumption of perfect substitution within schooling groups ($\sigma_e \rightarrow \infty$). However, some simple schooling models in both the U.S. (e.g., Card and Lemieux, 2001) and in a small subset of Latin American countries (Manacorda et al., 2010) allow for workers within schooling groups to be imperfect substitutes. If the simple schooling model above is properly-specified, so that actual aggregate production is approximated by equations (46) and (47), and if

potential experience groups are not perfectly-substitutable, equations (48) and (49) suggest that shifts in supply may help explain the patterns observed in Panels B and C of Figure 5.

First, the final RHS term of equation (48) suggests that if the (log) relative supply of more-educated workers among more-experienced men underwent a sharp trend increase in 2002, whereas changes over time in the (log) relative supply of more-educated workers among inexperienced men were relatively-smooth, this may explain differences over time in the log schooling premia observed in Panel B of Figure 5. However, figures from Table 12 imply that a story of this nature is unlikely to help, since increases over time in the relative supply of more-educated workers was smooth for both inexperience and experienced men: the annual rate of increase of the log relative supply of more-educated workers among inexperienced men was rapid and relatively-steady at roughly 12% per year from 1995-2002 and 9% per year from 2002-2011, and the annual rate of increase of the log relative supply of more-educated workers among experienced men was slower and steady at roughly 5% per year during both of these periods.

Second, can movements in supply explain the large decreases, within both schooling groups, in experience premia documented in Panel C of Figure 5? According to the last RHS term in equation (49), these decreases may be consistent with the simple schooling model if, within both schooling groups, the relative supply of more-experienced workers increased rapidly across the 1995-2011 earnings years, pushing down experience premia. However, Table 12 suggests that, while this may be true for non-upper-secondary-educated workers, for whom the log relative supply of more-experienced workers increased by roughly 4.4% and 4.9% annually for the 1995-2002 and 2002-2011 periods, respectively, this is not true for upper-secondary-educated workers, for whom the log relative supply of more-experienced workers *decreased* by roughly 2.8% annually from 1995-2002, then increased by only 0.8% from 2002-2011.

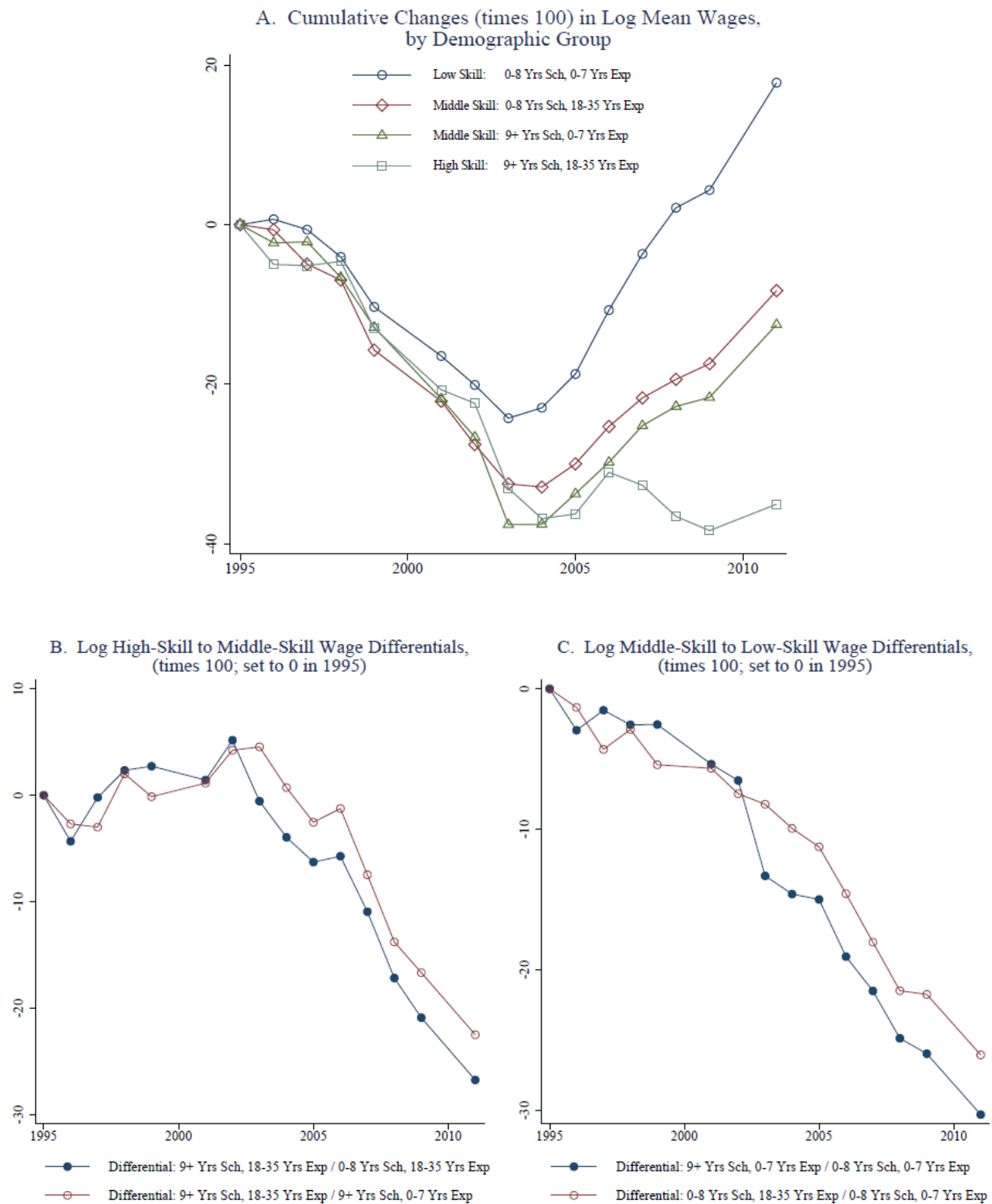
Re-Examining the Compression of Schooling Differentials, 2002-2011 Panels B and C of Figure 13 suggest that the simple model outlined in equations (46) and (47) appears inconsistent with the data. Moreover, Panel A of Figure 13 reveals a striking fact: the schooling premium, which held steady at roughly 2.5 from 1995-2002, then dropped to roughly 1.9 in 2011, has a strikingly-similar time-series to the experience premium, which held steady at roughly 2.05 from 1995-2002, then dropped to approximately 1.6 in 2011. Interestingly, this movement in the experience premium occurred concurrent with relatively-little movement of the relative supply of more-experienced work-

ers. Table 13 documents that the shares of employment represented by different potential experience groups changed very little, compared to movements across schooling groups, from 1995 to 2011: the employment share of inexperienced men, in efficiency units, decreased from 0.15 in 1995 to 0.13 from 1995, while the share of experienced men increased from 0.52 to 0.55. Taken together, the analysis of this section suggests that shifts in labor demand - potentially heterogeneous within schooling groups and strikingly-similar in their effects on the schooling and experience wage premia displayed in Panel A of Figure 13 - may play a role in the evolution of relative wages in Brazil from 1995-2011. In the next section, I develop a simple model that attempts to capture such changes.

II.D.2 A Simple Model of Low-, Middle- and High-Skill Workers

The previous section demonstrated that observed patterns of relative wages and supplies appear inconsistent with one standard model of the upper-secondary / non-upper-secondary wage premium. However, while simple schooling group models are particularly empirically-relevant for the U.S., the original supply and demand framework for examining changes in the relative wages of demographic groups of workers based on their individual characteristics, outlined in the seminal work of Katz and Murphy (1992), is quite general. In this section, I sketch a simple alternative to the model above that groups men into three skill groups based on their levels of both formal education and potential experience, and I use this model to briefly-examine the shifts in supply and demand potentially-responsible for changes in relative wages between these groups in Brazil from 1995-2011.

Figure 14: Wages of Low-Skill, Medium-Skill and High-Skill Male Workers, 1995-2011:



Two Motivating Facts The model is motivated by two facts. First, in contrast to the U.S. labor market, in which the Mincerian human capital earnings function is - to an approximation - separable in schooling and potential experience (Mincer, 1974; Murphy and Welch, 1990; Lemieux, 2006b), in Brazil potential-experience earnings profiles are decidedly-steeper among more-, relative to less-educated men (e.g., Leandro de Moura, 2008). For example, in 1995, while the composition-adjusted upper-secondary / non-upper-secondary wage ratio was 2 for men with 0-7 years of experience, it was 2.9 for men with 18-35 years of experience; similarly, while the composition-adjusted experience premium (the wage ratio of men with 18-35, relative to 0-7 years of experience) was 1.8 for men with 0-8 years of education, it was 2.6 for men with 9+ years of education.⁸⁵ As a result, the *level* of wages in Brazil is lowest for workers without schooling or experience, highest for those with both, and is decidedly more-similar among men with (i) less education and more experience, and (i) more education and less experience, compared to the U.S., in which the level of wages is divided more-distinctly along schooling lines.

Second, Figure 14 shows that this appears to also be true for *changes* in wages. In this figure, I examine the wages of three distinct skill groups. First, "low-skill" men are defined as those with neither advanced schooling or experience (0-8 years of schooling and 0-7 years of experience), "high-skill" men are defined as those with both advanced schooling and experience (9+ years of schooling and 18-35 years of experience), and "middle-skill" men are defined as those with one, but not both of those characteristics. Panel A of Figure 14 shows that changes over time in the log mean wage of the two displayed groups of middle-skill men - groups that contains both workers with upper-secondary education and non-upper-secondary education - are strikingly-similar.⁸⁶ Panels B and C of Figure 14 re-present the same wage differentials presented in Panels B and C of Figure 5, but instead of grouping them by schooling or experience, groups them by skill. Panels B and C of Figure 14 show that both log high-skill/middle-skill wage ratios are strikingly-similar, and feature an increase from 1995-2002, followed by rapid decline from 2002-2011, while the log middle-skill/low-skill wage ratios feature a more-steady slowdown across the entire 1995-2011 period. These patterns appear to match changes

⁸⁵In the U.S. from 1963-1987, by contrast, the level of the composition-adjusted wage ratio of college-, relative to high school-educated workers was nearly-identical for workers with 1-5 years of potential experience and all other workers, while the level of the composition-adjusted wage ratio of men with 26-35 / 1-5 years of potential experience was very similar for high school and college-educated workers (Katz and Murphy, 1992, Figure 1).

⁸⁶A simple regression of one time series on the other is associated with an R-squared statistic of 0.963.

observed in the overall wage distribution displayed in Panels A and B of Figure 4: while the top tail of the log wage distribution featured a sharper compression in 2002-2011 than in 1995-2011, the rate of compression among the extreme bottom tail of the log wage distribution was relatively more-steady across the entire 1995-2011 period.

A Supply and Demand Model In this section, I sketch a simple model that captures the idea that these three groups - low-, middle- and high-skilled workers - may be separate inputs in aggregate production. I follow the approach of Katz and Murphy in that I explicitly-model the non-comprehensive skill groups defined above, and then, in my empirical work, I assume that the supply of these groups can be captured by skill-"equivalents," constructed by combining the labor of all types of workers.⁸⁷ I assume that production of an aggregate output Y can be approximated by a CES function with three factors:

$$(52) \quad Y_t = \left[\sum_j (A_{jt} L_{jt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}, \quad j \in \{L, M, H\}$$

where L_t , M_t and H_t are indices that capture the effective units of labor supplied by low-, middle- and high-skill workers, respectively, A_{jt} is a labor-augmenting demand parameter associated with skill group j ,⁸⁸ and σ is a constant elasticity of substitution that represents the degree to which each of these three factors can be combined.⁸⁹ If the labor market is competitive, so that workers are paid their marginal products, relative wage equations take a simple form:

⁸⁷I discuss the construction of supply equivalents below.

⁸⁸The simplifying assumption that labor demand only takes labor-augmenting form (and not, for example, share form) does not affect the demand indices developed below.

⁸⁹For simplicity, equation (52) assumes that all pairs of skill groups are equally-substitutable; an alternative assumption, of course, is that elasticities may differ between different pairs of skill groups. For example, Goldin and Katz (2008; 2009) assume one constant elasticity of substitution between workers with some college versus all workers without any college, and a different elasticity of substitution between non-college workers with and without a high school diploma.

$$(53) \quad \ln \left(\frac{w_{Ht}}{w_{Mt}} \right) = \frac{\sigma - 1}{\sigma} \ln \left(\frac{A_{Ht}}{A_{Mt}} \right) - \frac{1}{\sigma} \ln \left(\frac{H_t}{M_t} \right)$$

$$(54) \quad \ln \left(\frac{w_{Mt}}{w_{Lt}} \right) = \frac{\sigma - 1}{\sigma} \ln \left(\frac{A_{Mt}}{A_{Lt}} \right) - \frac{1}{\sigma} \ln \left(\frac{M_t}{L_t} \right)$$

where w_{jt} represents the wage of workers in skill group j at time t . Equations (53) and (54) reflect the high-/middle-skill and middle-/low-skill wage ratios, respectively, and in each case show that change over time in these ratios is attributable either to shifts in demand (the 1st RHS terms) or shifts in supply (the 2nd RHS terms). Finally, equations (53) and (54) suggest two standard demand indices that capture the relative demand for high-/middle-skill and middle-/low-skill workers, respectively:⁹⁰

$$(55) \quad D_{Ht} = (\sigma - 1) \ln \left(\frac{A_{Ht}}{A_{Mt}} \right) = \sigma \ln \left(\frac{w_{Ht}}{w_{Mt}} \right) + \ln \left(\frac{H_t}{M_t} \right)$$

$$(56) \quad D_{Lt} = (\sigma - 1) \ln \left(\frac{A_{Mt}}{A_{Lt}} \right) = \sigma \ln \left(\frac{w_{Mt}}{w_{Lt}} \right) + \ln \left(\frac{M_t}{L_t} \right)$$

The quantity D_{Ht} in equation (55) captures the relative demand for high-, relative to middle-skill workers, and the quantity D_{Lt} in equation (56) captures the relative demand for middle-, relative to low-skill workers.

Constructing Relative Wage and Supply Series I follow Katz and Murphy in constructing the relative wage and supply series modeled in equations (53) and (54). For each relative wage ratio in the LHS of these equations, I use the log of the ratio of composition-adjusted mean wages for each group where, as discussed above, "low-skill" men are defined as those with neither advanced schooling or experience (0-8 years of schooling and 0-7 years of experience), "high-skill" men are defined as those with both advanced schooling and experience (9+ years of schooling and 18-35 years of experience), and "middle-skill" men are defined as those with one, but not both of those characteristics. For the relative supplies of low-, middle- and high-skill labor featured on the RHS of equations (53) and (54),

⁹⁰These indices are analogous to those developed in Katz and Murphy (1992) and used to measure change over time in the relative demand for college-educated labor in Autor et al. (2008).

I take a two-step approach. First, men who fall into these three categories are denoted as "pure" low-, middle- and high-skill equivalents. Second, the labor supplied by the remaining men - men with 8-17 years of potential experience - is allocated using a regression methodology. I divide these men into two schooling groups (0-8 years of schooling and 9+ years of schooling), and I follow Katz and Murphy by regressing the time series of log mean wages of these two groups onto the time series of log mean wages for the low-, middle- and high-skill workers (without a constant). The coefficients on this regression determine the shares of these groups, in efficiency units, allocated into each of the three skill bins. Reassuringly, both groups are primarily-allocated to the middle-skill group: for men with 8-17 years of potential experience and 0-8 years of education, allocation shares are 0.32 (low), 0.62 (middle) and -0.04 (high), while for men with 8-17 years of potential experience and 9+ years of education, allocation shares are -0.15 (low), 0.62 (middle) and 0.59 (high).⁹¹ Changes over time in the employment shares, in efficiency units, of the resulting low-, middle- and high-skill male supply equivalents are displayed in Figure 7.

⁹¹Under the methodology of Katz and Murphy (1992), these shares needn't add up to one or be weakly greater than zero.

TABLE 14
Annual Changes (times 100) in Log Relative Wages, Supplies,
and Demand Indices: 1995-2011

Comparison	1995-2002	2002-2011	1995-2011
<u>High-Skill / Middle-Skill:</u>			
Log Relative Wage	0.7	-3.4	-1.6
Log Relative Supply	3.3	3.3	3.3
Relative Demand Indices:			
$\sigma = 1$	4.0	-0.1	1.7
$\sigma = 2$	4.7	-3.5	0.1
$\sigma = 3$	5.4	-6.9	-1.5
<u>Middle-Skill / Low-Skill:</u>			
Log Relative Wage	-1.0	-2.2	-1.7
Log Relative Supply	6.3	16.3	11.9
Relative Demand Indices:			
$\sigma = 1$	5.2	14.1	10.2
$\sigma = 2$	4.2	11.9	8.5
$\sigma = 3$	3.2	9.7	6.8

Figure 15: Employment Shares, in Efficiency Units, of Low-Skill, Middle-Skill and High-Skill Male Equivalent Workers: 1995-2011

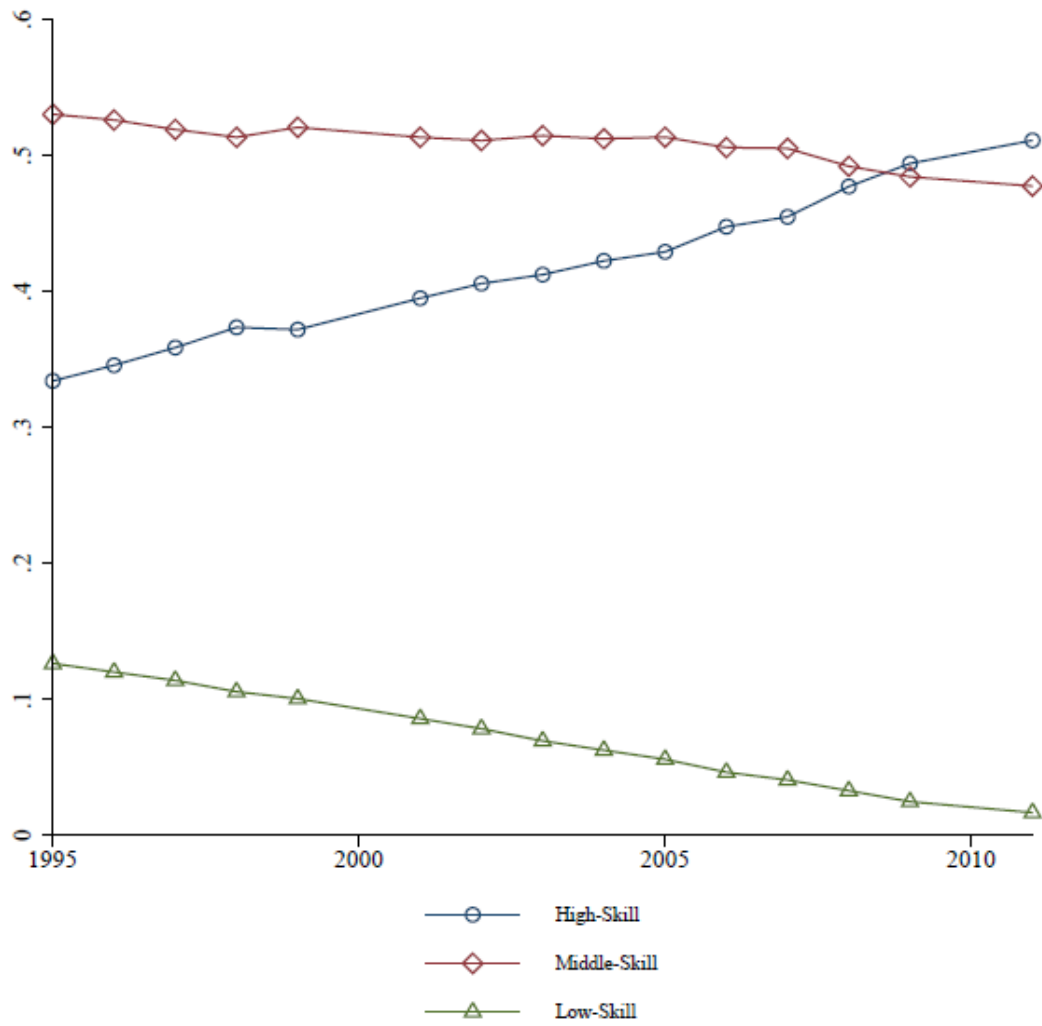
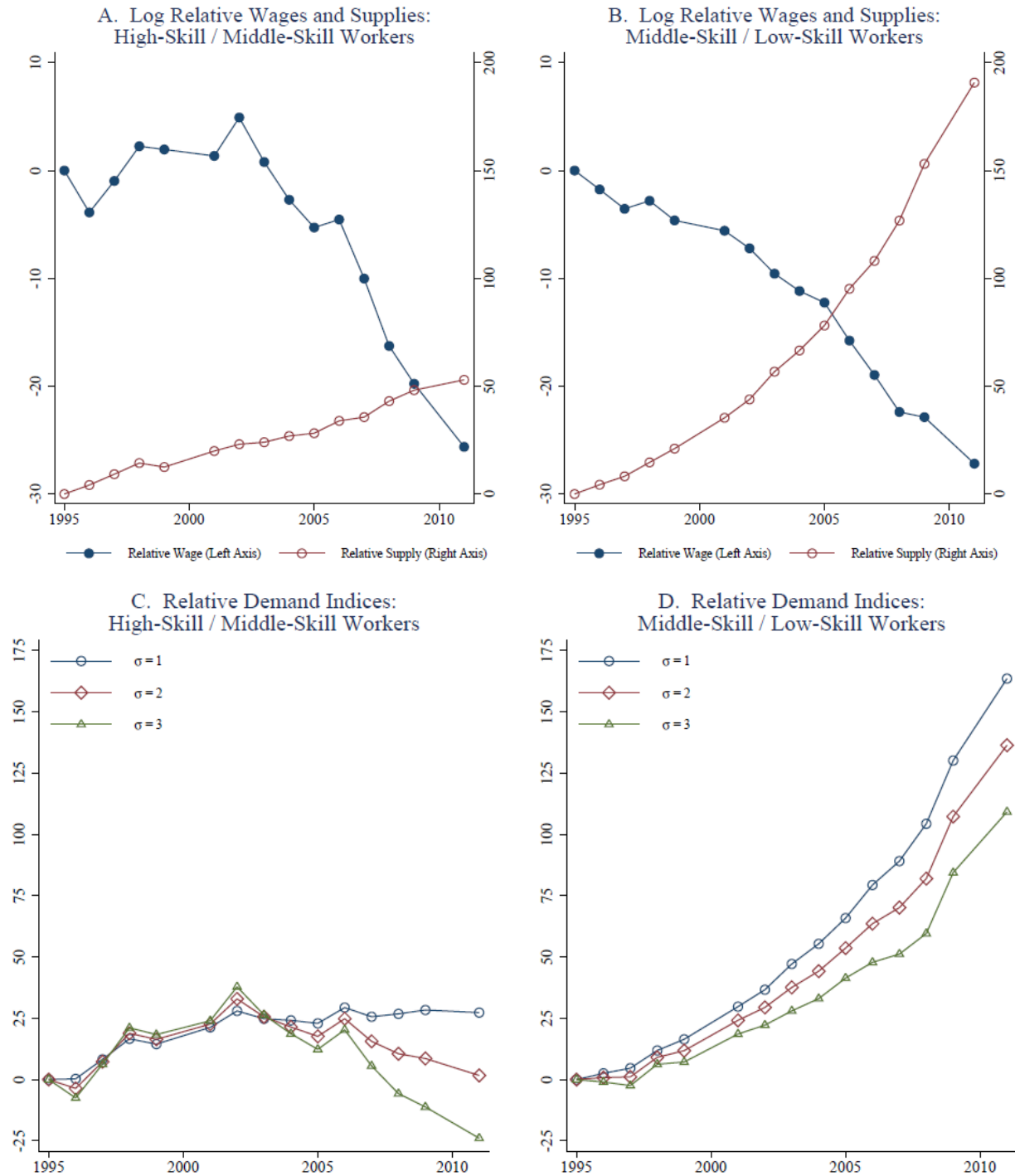


Figure 16: Low-Skill, Middle-Skill and High-Skill Workers: Relative Wages, Supplies and Demands (times 100; set to 0 in 1995): 1995-2011



Estimated Changes in the Supply and Demand for Skill Figure 16 and Table 14 summarize the analysis of high-/middle-skill and middle-/low-skill wage differentials using the simple framework described above. The relative wage series plotted in Panels A and B of Figure 16 confirm that while the high-/middle-skill wage differential features a sharp downward trend-break in 2002, change over time in the middle-/low-skill wage differential is much smoother. Figures in Table 14 confirm this: annual growth in the high-/middle-skill (middle-/low-skill) log relative wage series is roughly 0.7% (-1%) from 1995-2002, compared to -3.4% (-2.2%) from 2002-2011. However, while Panel A of Figure 16 shows steady and relatively-slow growth in the relative supply of high-/middle-skill workers (roughly 3.3% across the entire 1995-2011 period), Panel B shows that rate of growth of the relative supply of middle-/low-skill workers is rapidly increasing; figures from Table 14 reveal that annual growth in relative supply increased from 6.3% across the 1995-2002 earnings years to 16.3% across the 2002-2011 period. Inspection of Figure 15 shows that a driving force behind this rapid increase in relative supply is that, due to the expansion of education in Brazil, the share of young men (in efficiency units) without at least some upper-secondary education decreased almost to 0 in 2011.

These observed series of changes in relative wages and supplies suggest - for a wide range of plausible elasticity values considered in Table 14 and Panels C and D of Figure 16 - that, consistent with earlier work, significant trend changes in relative demand did indeed occur between the 1995-2002 and 2002-2011 periods. First, consider the relative demand for high-, relative to middle-skilled workers, defined in equation (55). Smooth growth in relative supply, combined with a trend break in relative wages beginning in 2002 favoring the middle-skilled suggests, for a range of values of σ from 1 to 3, that annual growth in the relative demand for high-skill workers was roughly 4-12% lower during the 2002-2011 earnings years compared to the previous seven years. Further, these changes in demand are not smooth: Panel C of Figure 16 reveals a sharp trend break in relative demand favoring middle-skilled workers in 2002 that roughly coincides with the beginning of the boom in the Brazilian economy described in Panel D of Figure 4.

By contrast, changes in demand for middle-, relative to low-skilled workers, appears not to contain a sharp decline in 2002, but instead to feature a smooth and convex increase across the entire 1995-2011 period. Panel B of Figure 16 shows that smooth and increasingly-rapid declines in the middle-, to low-skill relative wage ratio occurred despite tremendous and convex increases in the relative supply of middle-skilled workers. These patterns are potentially-reconciled by increases in relative demand that are large enough to keep the relative wage premium from decreasing even more-rapidly than observed.

Figures in Table 14 suggest that, for a range of values of σ from 1 to 3, annual growth in the relative demand for middle-skill workers, relative to low-skill workers, was roughly 6-9% faster from 2002-2011 than during the 1995-2002 earnings years.

Can Changes in Cohort Quality or Mis-Specification Explain the Results? The estimation of shifts in relative demand displayed in Table 14 and Panels C and D of Figure 16 depend heavily on at least two assumptions. First, that changes over time in relative wages reflect only changes in price, not changes in cohort quality. While one standard method of distinguishing between changes in the effects of cohort quality and changes in price is to track wages of the same cohort over time (e.g., Juhn et al., 1993 use this technique to argue that changes in the composition of unobserved skills are unlikely to explain a large fraction of increasing residual wage inequality in the United States), this method is complicated by the fact that the wage premia of interest in the current study depend on potential experience itself, which cannot be held fixed over time within the same cohort. A second assumption underlying the results of this section is that the model is properly-specified, so that equation (52) approximates the true structure of aggregate production in Brazil. With respect to the estimated shifts in relative demand for middle-, relative to low-skill workers outlined, for example, in Panel D of Figure 8, one potentially-troubling aspect in this regard is that the near-complete disappearance of young men without some upper-secondary education across the 1995-2011 earnings years (as seen in Figure 7) may have led to an over-estimate of increases in the relative supply of middle-skilled, relative to low-skilled workers. Because both the assumption of fixed quality within cohort and proper model specification are difficult to verify, I consider the estimated relative demand indices displayed in Table 14 and Panels C and D of Figure 16 to be suggestive.

II.E Conclusion

Patterns of schooling premia in Brazil - stable during the relative stagnation of 1995-2002, then dramatically compressing during the 2002-2011 economic boom - combined with a smooth schooling expansion seem to suggest a simple supply and demand story: that relative labor demand shifted towards less-educated workers in Brazil, concurrent with the international increase in commodity prices beginning in 2002. In this paper, I argue that this simple story may not be consistent with observed movements in relative wages and supplies across the 1995-2011 period. First, a simple measure of the schooling premium - the upper-secondary/non-upper-secondary wage ratio - is strikingly different

for less- versus more-experienced workers. Second, within each schooling group, experience premia decreased dramatically. As an alternative to the simple schooling group approach, I sketched a simple model in which low-, middle- and high-skill workers, groups defined by both schooling and potential experience, are assumed to be separate inputs in aggregate production. Application of this model to the 1995-2011 earnings years provides suggestive evidence that the economic boom of 2002-2011 may have been associated with shifts in relative demand toward middle-skilled workers, rather than only towards less-educated workers.

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IV Appendices

Appendix A.1: Description of Individual-Level Source Variables

In this appendix, I briefly describe the individual-level source variables I use from the cleaned 1964-2009 March CPS data files of Acemoglu and Autor (2011) described in the text.¹ Acemoglu and Autor's March CPS data files contain the following individual-level demographic source variables: a CPS sample weight, dummy for male, age at earnings year, years of schooling completed, 5 "narrow" schooling groups defined in Autor et al. (2008), years of potential labor market experience defined as in Autor et al. (2008) as the minimum of age minus years of schooling minus 7 and age minus 17, and a race variable that contains three values (white, black and other). The data files also contain individual-level variables on employment: a dummy variable that flags an individual as a full-time, full-year (FTFY) worker, weeks worked in the earnings year, the usual hours per week worked in the earnings year, weekly wage for wage and salary earners, and a PCE wage deflator.

¹These files are generously provided by the authors: <http://economics.mit.edu/faculty/dautor/data/acemoglu>.

Appendix A.2: Construction of the Empirical Measures $\widehat{\lambda}_{sge}$, $\widehat{\ln w_{sget}}$ and $\widehat{g_{set}}$

In this appendix, I define the empirical measures of employment shares $\widehat{\lambda}_{sge}$, wages $\widehat{\ln w_{sget}}$ and labor supply $\widehat{g_{set}}$ described in the text. Each of these measures captures quantities relevant for the 16 "narrow" schooling-by-gender-by-potential experience groups defined in the text.

Employment Shares $\widehat{\lambda}_{sge}$: The empirical measure $\widehat{\lambda}_{sge}$ is intended to capture the employment share, averaged over a base set of years, of workers in schooling group s , gender g and potential experience group e , relative to the total employment of all workers within schooling group s . I follow the empirical literature, beginning with Katz and Murphy (1992) in (i) using the simple share of raw hours worked by all employed workers (as measured in the quantity sample defined in the text) in each demographic group in year t , divided by total raw hours worked in that year, to generate a yearly employment share, and then (ii) averaging these shares across a set of years to produce the employment shares used to weight log wage measures. Formally, I define:

$$(57) \quad \widehat{\lambda}_{sge} = \frac{1}{T} \sum_{t=1}^T \left[\frac{\sum_{i \in sget} (\omega_{sgeit} \cdot h_{sgeit})}{\sum_{g'} \sum_{e'} \sum_{i' \in sg'e't} (\omega_{sg'e'it} \cdot h_{sg'e'it})} \right]$$

where ω_{sgeit} is the CPS individual sample weight and h_{sgeit} is the usual hours worked per week times the total weeks worked for worker i in schooling group s ,² of gender g , and in potential experience group e with reference to the earnings year t and where, with some abuse of notation, I define " $sget$ " as the set of all such workers i in the quantity sample. Equation (57) shows that $\widehat{\lambda}_{sge}$ is the result of averaging over a set of time periods: the leading case in my empirical work (and the standard set of time periods in the literature) is averaging over all available earnings years.

Wage Measures $\widehat{\ln w_{sget}}$: Here I define the quantity $\widehat{\ln w_{sget}}$, which is intended to capture the log wage of workers in schooling group s , gender g and potential experience e during earnings year t . First, consider the 5 schooling groups of Autor et al. (2008), which divide workers into those with 11 or fewer, exactly 12, 13-15, exactly 16, and 17+ years of schooling. In order to clearly distinguish my treatment of these schooling groups from the alternative grouping $s \in \{H, C\}$ described above,

²Here, as is standard, s refers to (non-exhaustive) schooling groups, not schooling group s *equivalents*. That is, in this context, $s = H$ refers only to workers who completed exactly 12 years of schooling and $s = C$ refers only to workers who completed 16 or more years of schooling before labor market entry.

I denote the 5 Autor et al. groups by the symbol \tilde{s} . I then create the wage measures $\overline{\ln w_{\tilde{s}get}}$ in each year for each of 40 \tilde{s} -by- g -by- e demographic groups. Following Autor et al., I regress log weekly wages of FTFY workers from the wage sample, separately for each year and sex, on dummy variables for 4 education categories, a quartic in potential experience, a dummy for "black" and a dummy for "other (non-white) race," and interactions of the potential experience quartic with 3 broad schooling categories (11 or fewer, 13-15, and 16 or more years of education). The measure $\overline{\ln w_{\tilde{s}get}}$ is the predicted log wage from these regressions for white workers with schooling level \tilde{s} and with the mean experience level within each group e (5, 15, 25 and 35 years of potential experience, respectively).

The empirical measure $\overline{\ln w_{Hget}}$ for a worker in groups g and e and for year t is simply the corresponding wage measure for workers with exactly 12 years of schooling. The empirical measure of the log wage of college-educated workers for a worker in groups g and e in year t is defined as the fixed-weighted sum:

$$(58) \quad \overline{\ln w_{Cget}} = \widehat{\nu}_{ge} \cdot \overline{\ln w_{17+get}} + (1 - \widehat{\nu}_{ge}) \cdot \overline{\ln w_{16get}}$$

where $\overline{\ln w_{17+get}}$ and $\overline{\ln w_{16get}}$ are the wage measures for groups g and e at time t for groups representing workers with 17 or more and exactly 16 years of education, respectively, and where $\widehat{\nu}_{ge}$ is defined as the share of raw hours worked by workers with 17 or more years of education among college-educated workers of gender g and potential experience e , averaged across a set of base years, as measured in the quantity sample defined in the text. Formally:

$$(59) \quad \widehat{\nu}_{ge} = \frac{1}{T} \sum_{t=1}^T \left[\frac{\sum_{i \in 17+get} (\omega_{17+geit} \cdot h_{17+geit})}{\sum_{i \in 17+get} (\omega_{17+geit} \cdot h_{17+geit}) + \sum_{i \in 16get} (\omega_{16geit} \cdot h_{16geit})} \right]$$

where for $\tilde{s} \in \{17+, 16\}$, $\omega_{\tilde{s}geit}$, $h_{\tilde{s}geit}$ and the set $\tilde{s}get$ are defined similarly to ω_{sgeit} , h_{sgeit} and the set $sget$ above.

Labor Supply Measures $\widehat{g_{set}}$: The empirical measure $\widehat{g_{set}}$ is intended to capture the quantity of labor supplied by schooling group s equivalent workers of gender g and potential experience e at time t . I follow the empirical literature, beginning with Katz and Murphy (1992), in combining the labor supplied by all schooling groups into measures of high school and college "equivalent" labor. In

particular, for workers of gender g and potential experience e , I combine workers with 11 or fewer, exactly 12 and 13-15 years of schooling to form the measure of high school equivalent labor $\widehat{g_{Het}}$, and I combine workers with exactly 16, 17+ and 13-15 years of education to form the measure of college equivalent labor $\widehat{g_{Cet}}$.

In order to create these labor supply measures, I perform two steps similar to those in Autor et al. (2008). First, for any g and e , I define high school equivalent labor as the sum of labor supplied by workers with 11 or fewer and exactly 12 years of education plus 0.5 times the labor supplied by workers with 13-15 years of education, and I define college equivalent labor as the sum of labor supplied by workers with exactly 16 and 17+ years of education plus 0.5 times the labor supplied by workers with 13-15 years of education. Second, within each g , e and s , I weight raw hours worked by mean (non-log) wage, relative to the efficiency weight of the group. Formally, for each s , g and e , I construct the estimate $\widehat{\pi_{sge}}^{\text{p.s.}}$ by estimating equation (22) under the assumption of perfect substitution between experience groups ($\sigma_e \rightarrow \infty$).³ I define:

$$(60) \quad \widehat{g_{Het}} = \left(\frac{\widehat{\gamma_{11-ge}}}{\widehat{\pi_{Hge}}^{\text{p.s.}}} \right) \cdot l_{11-ge} + \left(\frac{\widehat{\gamma_{12ge}}}{\widehat{\pi_{Hge}}^{\text{p.s.}}} \right) \cdot l_{12ge} + \left(\frac{\widehat{\gamma_{13-15ge}}}{\widehat{\pi_{Hge}}^{\text{p.s.}}} \right) \cdot \frac{l_{13-15ge}}{2}$$

$$(61) \quad \widehat{g_{Cet}} = \left(\frac{\widehat{\gamma_{16ge}}}{\widehat{\pi_{Cge}}^{\text{p.s.}}} \right) \cdot l_{11-ge} + \left(\frac{\widehat{\gamma_{17+ge}}}{\widehat{\pi_{Cge}}^{\text{p.s.}}} \right) \cdot l_{12ge} + \left(\frac{\widehat{\gamma_{13-15ge}}}{\widehat{\pi_{Cge}}^{\text{p.s.}}} \right) \cdot \frac{l_{13-15ge}}{2}$$

where for the demographic group defined by schooling group \tilde{s} , gender g and potential experience group e , $\widehat{\gamma_{\tilde{s}ge}}$ is the mean (non-log) wage of workers (from the wage sample), relative to a base group of males with exactly 12 years of schooling and 10-19 years of potential experience, across the period $t = 1, \dots, T$ defined above, and $l_{\tilde{s}get}$ is the total raw hours supplied by the group (from the quantity sample) during earnings year t .

³In equation (18), I use the supply measures $\widehat{g_{set}}$ in estimating σ_e : thus the estimated value of σ_e from this equation cannot be used to generate this supply measure. Given no a-priori information about σ_e , a natural step is to assume perfect substitution between potential experience groups in equation (22). This procedure allows me to estimate each π_{sge} without information on the supply measures $\widehat{g_{set}}$.

Appendix A.3: Derivation of Equations (11), (12) and (13)

In this appendix, I derive equations (11), (12) and (13) in the text. This derivation (i) begins with a general CES aggregate in order to emphasize that the relative demand measures \widehat{D}_t , \widehat{D}_{st} and \widehat{D}_{sget} in the text are each special cases of the same general relative demand measure, and (ii) allows technology within this CES aggregate to take both share form and labor-augmenting form in order to emphasize that these relative demand measures do not depend on the simplifying assumption in equations (52), (2) and (3) that technology takes labor-augmenting form.

Consider a CES aggregate Q of the form:

$$(62) \quad Q_t = \left[\sum_{k=1}^K \alpha_{kt} (A_{kt} N_{kt})^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}}$$

where Q_t is the value of the aggregate in year t , k indexes K sub-aggregates within Q , α_{kt} and A_{kt} are sub-aggregate k -specific technology parameters of the share form and of the labor-augmenting form, respectively, N_{kt} denotes the k th sub-aggregate and σ represents the elasticity of substitution between each sub-aggregate. Denote the partial derivative $\partial Q_t / \partial N_{kt}$ by φ_{kt} . By definition:

$$(63) \quad \varphi_{kt} = \left(\frac{Q_t}{N_{kt}} \right)^{\frac{1}{\sigma}} \alpha_{kt} A_{kt}^{\frac{\sigma-1}{\sigma}}$$

Equation (63) suggests a comparison between k and all other sub-aggregates within Q_t :

$$(64) \quad \widetilde{\varphi}_{kt} = \widetilde{\alpha}_{kt} + \frac{\sigma-1}{\sigma} \cdot \widetilde{A}_{kt} - \frac{1}{\sigma} \widetilde{N}_{kt}$$

where

$$\widetilde{\varphi}_{kt} = \ln \varphi_{kt} - \left(\frac{1}{K-1} \right) \sum_{k' \neq k} \ln \varphi_{k't}$$

and the terms $\widetilde{\alpha}_{kt}$, \widetilde{A}_{kt} and \widetilde{N}_{kt} are defined similarly. I can re-arrange equation (64) to yield:

$$(65) \quad \sigma \widetilde{\alpha_{kt}} + (\sigma - 1) \widetilde{A_{kt}} = \sigma \widetilde{\varphi_{kt}} + \widetilde{N_{kt}}$$

Equation (65) immediately yields equations (11), (12) and (13) in the text. To see this, note that equation (65) is equivalent to equation (11) for the case that $Q_t = Y_t$ for all t , $\alpha_{kt} = 1$ for all k and t , $A_{kt} = A_{st} = \widehat{A_{st}}$ for $s \in \{C, H\}$ and for all t , $\sigma = \sigma_s = \widehat{\sigma_s}$, $k = C$, $k' = H$, $K = 2$, $N_{Ct} = C_t$ and $N_{Ht} = H_t$ (with some abuse of notation), and the labor market is competitive so that $\widetilde{\varphi_{Ct}} = \ln(w_{Ct}/w_{Ht})$ where w_{st} represents the marginal product of the sub-aggregate s at time t . Similar steps establish that equation (65) is equivalent to equations (12) and (13).

Appendix A.4: Derivation of Equation (21)

In this appendix, I derive equation (21) in the text, which represents the bias of the estimate of σ_e implied from a "short" variant of estimating equation (18) in which $rel.demand_{get}$ is omitted.

Suppose that the error term ε_{get} in equation (18) is uncorrelated with $rel.supply_{get}$ for all gender-by-potential experience groups. Then, consider running the "short" regression:

$$(66) \quad rel.wage_{get} = \psi_{ge} - \frac{1}{\sigma_e^{\text{short}}} \cdot rel.supply_{get} + \zeta_{get}$$

separately for each group. Here I have implicitly defined the regression coefficient of interest as $-1/\sigma_e^{\text{short}}$ to emphasize that regression (66) is intended to measure the structural parameter σ_e , and I have defined the regression constant and error term, ψ_{ge} and ζ_{get} respectively, to emphasize the distinction between the theoretical and empirical parameters. By the definition of a regression coefficient:

$$\sigma_e^{\text{short}} = - \left[\frac{Cov(rel.wage_{get}, rel.supply_{get})}{Var(rel.supply_{get})} \right]^{-1}$$

substituting for $rel.wage_{get}$ using equation (18) yields:

$$\sigma_e^{\text{short}} = \sigma_e - \frac{\sigma_e}{\sigma_e - 1} \cdot \left[\frac{Cov(rel.demand_{get}, rel.supply_{get})}{Var(rel.supply_{get})} \right]^{-1}$$

which, after defining $\widehat{\sigma}_e^{\text{short}}$ as the OLS estimate of σ_e from the short regression, yields the bias in equation (21).

Appendix A.5: Two-Step Estimation of σ_e using Equations (18) and (20)

In this appendix, I show that if changes in relative demand are accurately-captured by equation (20), the two-step procedure outlined in the text for estimating σ_e in pooled regressions using equations (18) and (20) yields consistent estimates of σ_e . First, consider the case where $f_{ge}(\cdot)$ is a quadratic polynomial. In the first step, I run the regressions:

$$(67) \quad \begin{aligned} \widehat{rel.wage}_{get} &= \varphi_{0wge} + \varphi_{1wge}t + \varphi_{2wge}t^2 + \widetilde{w}_{get} \\ \widehat{rel.supply}_{get} &= \varphi_{0sge} + \varphi_{1sge}t + \varphi_{2sge}t^2 + \widetilde{s}_{get} \end{aligned}$$

separately for each gender-by-potential experience group. The residuals from these regressions, \widetilde{w}_{get} and \widetilde{s}_{get} , are those components of relative wages and supplies that are orthogonal, by definition, to the ge -specific quadratic function of time. They are then used in the second step regression:

$$(68) \quad \widetilde{w}_{get} = \varphi_{ge} + \rho \cdot \widetilde{s}_{get} + e_{get}$$

The second step regression coefficient has the form:

$$(69) \quad \rho = \frac{Cov(\widetilde{w}_{get}, \widetilde{s}_{get})}{Var(\widetilde{s}_{get})} = \frac{Cov(\widehat{rel.wage}_{get}, \widetilde{s}_{get})}{Var(\widetilde{s}_{get})} = -\frac{1}{\sigma_e}$$

The first and last equalities in expression (69) result from the definition of regression coefficients, applied to equations (68) and (18) and (20), respectively, and from assuming that the theoretical relative wage and supply terms equal their empirical counterparts. The middle equality in expression (69) results from plugging in the expression for the relative wage from equation (67):

$$Cov(\widetilde{w}_{get}, \widetilde{s}_{get}) = Cov(\widehat{rel.wage}_{get} - \varphi_{0wge} - \varphi_{1wge}t - \varphi_{2wge}t^2, \widetilde{s}_{get}) = Cov(\text{rel. wage}_{get}, \widetilde{s}_{get})$$

where the second equality results from the fact that the residual \widetilde{s}_{get} is mechanically orthogonal to

t and t^2 . Steps analogous to these yield an expression similar to equation (69) for the case in which $f_{ge}(\cdot)$ is linear.

Appendix A.6: Log Mean Wage Decomposition Displayed in Table 12

In this appendix, I describe the decomposition used in Table 12. Denote an individual group by i and a broader group by g . In Panel A of Table 12, I display changes in the simple log mean wage of a given broad group g between times t_0 and t_1 :

$$\ln \bar{w}_{g,t_1} - \ln \bar{w}_{g,t_0}$$

The decomposition presented in Table 12 is:

$$\ln \bar{w}_{g,t_1} - \ln \bar{w}_{g,t_0} = \underbrace{\sum_{i \in G} \frac{N_i}{N_g} (\ln \bar{w}_{i,t_1} - \ln \bar{w}_{i,t_0})}_{\text{"composition-adjusted" component}} + \underbrace{\sum_{i \in G} \frac{N_i}{N_g} [(\ln \bar{w}_{g,t_1} - \ln \bar{w}_{g,t_0}) - (\ln \bar{w}_{i,t_1} - \ln \bar{w}_{i,t_0})]}_{\text{"composition effect" component}}$$

where N_i and N_g are the the average shares of total hours worked across the entire sample period by workers in cell i and group g , respectively, $\bar{w}_{i,t}$ is the mean of w among of workers in cell i in year t , and G is the set of all cells that compose group g . The first term on the RHS (and displayed in Panel B of Table 12) is the quantity of interest, and represents the change in the log mean had composition, in terms of total hours worked of sub-groups i , been held fixed within group g . The second term, displayed in Panel C of Table 12, reflects the fact that total hours worked within group g shifted between sub-groups.