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Abstract

We document the growth in higher education costs and tuition over the past 50 years. To explain these trends, we develop a general equilibrium model with skill- and sector-biased technical change. Finding the model's parameters through a combination of estimation and calibration, we show that it can explain the rise in college costs between 1961 and 2009, along with the increase in college attainment and the change in the relative earnings of college graduates. The model predicts that if college costs had ceased to grow after 1961, enrollment in 2010 would be 3 to 6 percent higher.

Keywords: higher education; skill-biased technical change

JEL classification numbers: I23; J24; D58

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1 Introduction

For most American families, the cost of a college education is a significant expense. College tuition has grown faster than inflation for decades. Many observers believe that, in the words of a recent CNN headline (Censky, 2011), "surging college costs price out [the] middle class". The Obama administration has proposed several measures to make college more affordable (White House, 2012). In this paper, we develop a general equilibrium model that explains why college costs have risen so dramatically, and consider its implications.

We begin by documenting trends in higher education costs and tuition over the past 50 years. The data show that the total cost of educating a college student has risen at roughly the same rate as per capita GDP. Since 1960, listed or "sticker price" tuition has grown more quickly than GDP, while tuition net of grant aid has risen at the same rate as GDP.

To explain the cost trend, we construct and analyze a general equilibrium model with skill- and sector-biased technical change. Having a general equilibrium model allows us to disentangle the interactions between the labor market and the higher education sector. In our model, forward-looking individuals make education decisions based on labor market returns and the cost of tuition. The higher education sector produces skilled workers using capital and skilled workers as inputs, and suffers from the service sector disease (Baumol and Bowen, 1966; Baumol, 1967; Baumol, 2012). In particular, we assume that total factor productivity (TFP) in the higher education sector is constant, and that colleges have only a limited ability to replace labor inputs with capital. The data appear consistent with such an assumption. For example, in 1976 there were 0.185 employees for each college student; in 2009 the number was essentially unchanged, at 0.186.[1] As college professors and administrators become more productive in other sectors, their wages, and the cost of college education, will rise.

We estimate the parameters of the model to match the observed time paths of educational attainment, the college earnings premium, higher education expenditures, higher education capital inputs, and GDP per capita over the past five decades. Our model successfully replicates the dramatic increase in higher education costs between 1961 and 2009. We then use the model to perform a number of numerical experiments. We find that the increase in enrollment is almost entirely due to skill-biased technological change in the non-education sector. Skill-biased change also causes college costs to grow more

[1] See Table 1 below.

quickly than GDP. In contrast, skill-neutral change (TFP growth) in the non-education sector causes college costs to grow more slowly than GDP. Nonetheless, TFP growth accounts for the majority of the cost increase, because it accounts for the majority of wage growth. We then use the model to assess the impact of the rise in higher education costs. Numerical experiments suggest that if college costs were at their 1961 levels, enrollment in 2010 would be 1 to 2 percentage points (3 to 6 percent) higher. We similarly find that the effects of changing tuition subsidies are small. Depending on the duration of the tuition change, we find enrollment-price elasticities of 0.04 to 0.07. These effects are modest, but consistent with a number of previous studies.

Our paper straddles two areas of research. The first is the industry-level analysis of higher education costs and tuition. The service sector disease is just one of several explanations for the increase in college costs and tuition. Many analysts argue that institutions of higher education have become increasingly inefficient. The inefficiencies arise from market power and public subsidies that allow colleges to pad their expenses (e.g., Bowen, 1980; Martin and Hill, 2013), or costly "arms races" (Ehrenberg, 2002). Other explanations focus on the tuition colleges charge, rather than the costs they incur.[2] In their recent book, Archibald and Feldman (2011) review the competing explanations and conclude that the service sector disease plays a central role. They show that the cost trajectory of higher education resembles those of other high-skill services, and that costs have risen rapidly at community colleges as well as at Ivy-league institutions.

Our contribution is to show that in a quantitative general equilibrium model, this combination of skill- and sector-biased technical change generates an increase in college costs similar to those actually observed. Moreover, the cost increases generated by this model are consistent with observed changes in educational capital, educational attainment, and earnings.

In addition to imposing external discipline on our estimates of the higher education cost function, using a quantitative general equilibrium model allows us to decompose the cost increase and perform a number of policy experiments. Our model thus contributes to a second area of research, the general equilibrium analysis of human capital accumulation and earnings dynamics. Ljungqvist (1993) argues that because the cost of providing education depends on the cost of skilled labor, education may be particularly expen-

[2]Many institutions, especially private ones, post a "sticker price" well in excess of the discounted tuition most students actually pay (Hoxby, 1997; Fu, 2010). Increasing the sticker price more quickly than net tuition has allowed colleges to offer a broader menu of net prices, increasing their ability to price discriminate. Another argument is that decreased public funding has forced schools to raise both gross and net tuition.

sive in countries where the current level of educational attainment is low. In contrast to Ljungqvist (1993), who proceeds theoretically, most of this literature is quantitative. In an influential paper, Heckman, Lochner and Taber (1998a) show that sustained skill-biased technological change can explain the changes in education and earnings observed over the past few decades. Lee and Wolpin (2006, 2010) and He and Liu (2008) consider similar topics. Akyol and Athreya (2005) emphasize that investments in college are risky: drop-out rates are high. Abbot, Gallipoli, Meghir and Violante (2013) analyze tuition schedules that vary with income and ability. Restuccia and Vandenbroucke (2013) and Castro and Coen-Pirani (2012) study educational attainment over much of the 20th century.

Our contribution is to determine the cost of higher education within the model, allowing it to adjust to economic events and affect education decisions. The paper with the cost structure most similar to ours is Garriga and Keightley (2007), who assume the cost of college is proportional to the wage for skilled labor. Garriga and Keightly, however, analyze stationary equilibria, while we consider transitions, and they focus on education policy rather than costs.[3]

The rest of the paper is organized as follows. In Section 2, we summarize historical patterns of higher education costs and pricing. In Section 3, we describe our model. In Section 4 we discuss how we estimate and/or calibrate the model's parameters, and assess how well the model fits the data. In Section 5, we use the model to perform policy analyses. In Section 6, we discuss some extensions to the model and conclude.

2 Higher Education Data

In this section, we present data related to higher education: costs, prices, enrollment, and earnings. We use these data to motivate the structure of our model, and as targets for estimating our model's parameters. A detailed description of how we constructed all of our data can be found in Appendix A.

2.1 Expenditures and Tuition

In considering "college costs", it is useful to distinguish between three distinct objects: (1) expenditures, the costs incurred in educating students; (2) the listed or "sticker" price for tuition; and (3) the net tuition students pay after receiving financial aid.

[3]In earlier drafts, Castro and Coen-Pirani (2012) also used a cost structure similar to ours. Their focus, however, was on educational attainment, rather than college costs.

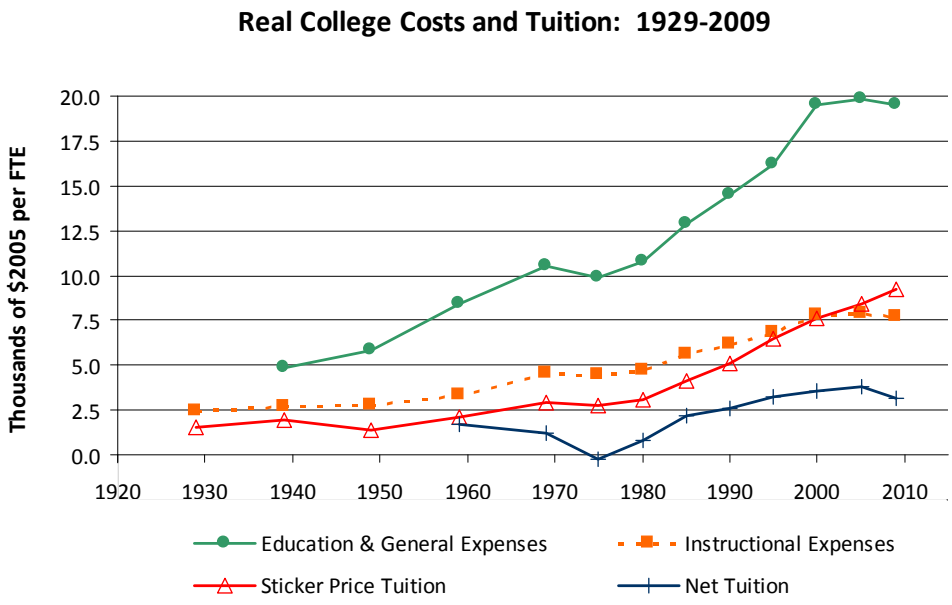

Figure 1

The top line in Figure 1 shows real expenditures per full-time-equivalent (FTE) student for the entire higher education sector, including 2-year, 4-year and graduate students. These data are drawn from the U.S. Department of Education's Digest of Education Statistics, and converted to 2005 dollars with the GDP deflator. The data are measured over academic years, July 1 - June 30, which we index by the initial calendar year. We focus on the subset of costs included in the education and general (E&G) category. Among the costs excluded from this category are auxiliary operations such as dormitories and hospitals. Average expenditures have quadrupled over time, from under $5,000 in 1939 to almost $20,000 today. Figure 1 also shows the "instructional" component of education and general expenditures. While we view any attempt to allocate costs in detail as problematic – for example, graduate research facilities are needed to attract faculty at top undergraduate institutions – it bears noting that instructional expenditures follow a pattern very similar to that of education and general costs.

Figure 1 shows sticker price tuition per FTE over the same time period. Consistent with the cost measures, we calculate sticker price tuition as tuition revenues divided by FTE, again using data from the Digest of Education Statistics. After falling between 1939 and 1949, sticker price tuition has risen more rapidly than expenditures. Most college students, however, receive some form of grant aid, either from the institution itself, or some external source such as the Federal government. Combining data from the Digest

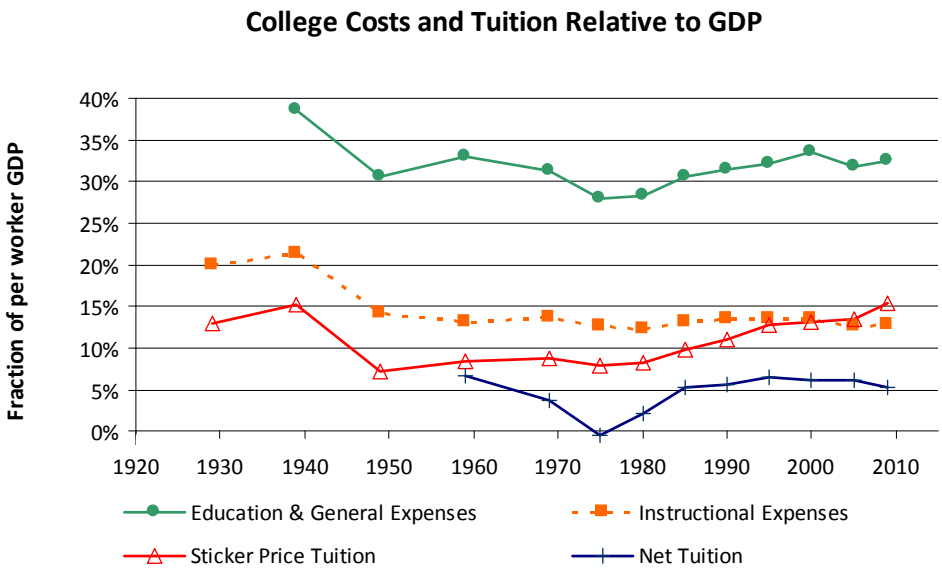

Figure 2

of Education Statistics and the College Board (2011), we calculate average student grant aid.[4] Subtracting aid from sticker price tuition yields net tuition. The bottom line in Figure 1 shows net tuition. Net tuition has grown more slowly over time than sticker prices, suggesting that the increase in sticker prices has been at least partly intended to increase the scope for price discrimination (Hoxby, 1997; Fu, 2010). Net tuition was particularly low in the 1970s, when many veterans received assistance. A striking feature of sticker price and especially net tuition is that they are quite low relative to expenditures. Although state aid to public institutions has not increased over time (in real per FTE terms), it is still significant. Federal grants are also a major source of income, even at private institutions.

Figure 2 shows the same costs and tuition as fractions of per capita GDP, which is constructed from the national accounts and Census data.[5] These data show that, with the exception of 1939, education and general expenditures have stayed between 27 and 34 percent of per capita GDP. After falling sharply between 1929 and 1949, sticker price tuition has grown faster than GDP, while net tuition has grown at the same rate as GDP.

[4] Our calculations omit tax benefits. In recent years, the Federal income tax credits for higher education have grown rapidly (College Board, 2011).

[5] Consistent with the model below, our measure of population consists of people aged 18-74. Calculating per capita GDP using the total population makes the series in Figure 2 more variable, but does not change their general properties. To be consistent with academic years, we use averages of consecutive calendar-year values.

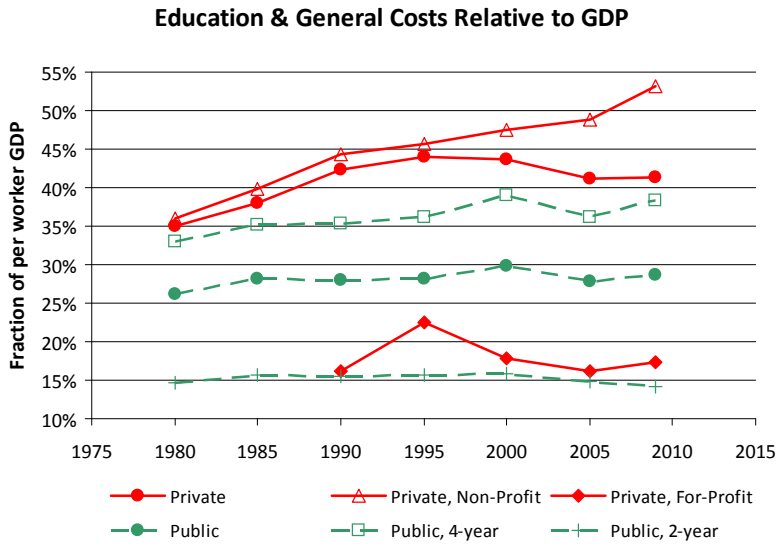

Figure 3

The net tuition ratio, being a function of averages, need not imply that higher education is as "affordable" today as it was in the past. Median income has grown far more slowly than per capita GDP, and tuition varies widely across students (Leonhardt, 2009).

The data in Figures 1 and 2 are averages taken across both public and private institutions, with the latter including for-profit institutions as well as traditional non-profit schools. These averages also combine data for 2-year and 4-year institutions. Figure 3 shows disaggregated expenditure data, expressed as a fraction of GDP. Perhaps the most notable feature of the data is the steady rise in expenditures at private not-for-profit institutions. Such an increase is consistent with the argument that elite private institutions are engaging in expensive "arms races". Figure 3 also shows expenditures for all private institutions. In the recent decade for-profit institutions have grown rapidly, from 4 percent of private enrollment in 1980 to 33 percent in 2009. Because for-profit institutions tend to have lower costs, their growth has pulled down average expenditures in the private sector. Although separate cost data for 2-year and 4-year institutions are available before 1996 only for public institutions, about 90 percent of 2-year students are in public colleges. Figure 3 reveals that once 2-year public institutions are excluded, public and private institutions have fairly similar levels of expenditure. The data targets to which we fit our model will exclude the 2-year public institutions: we will refer to this group as "4-year" institutions. Figure 3 also shows, however, that expenditures at 2-year public colleges follow a trajectory similar to that at 4-year public institutions; Archibald

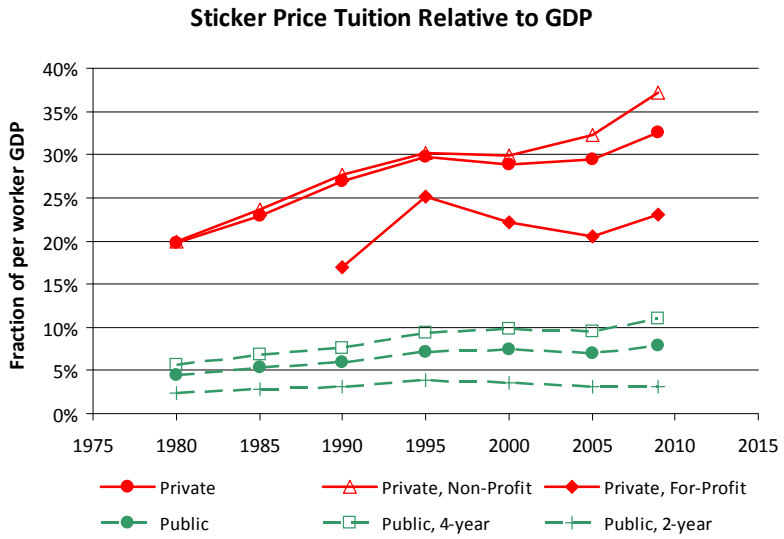

Figure 4

and Feldman (2011) view this as evidence against the arms race hypothesis.

Figure 4 displays disaggregated sticker price tuition. The distinction between non-profit and for-profit private institutions is relatively modest. Although for-profit institutions have considerably lower costs, they rely more heavily on tuition revenue. Disaggregated revenue data are first available in 1992; we use the 1992 tuition ratio to infer 4-year tuition for earlier years. Since 1992, tuition for 4-year public institutions has risen relative to tuition for 2-year public institutions.

2.2 Staffing and Compensation

Table 1 shows staffing levels, measured as the ratio of students (in FTE) to employees (in FTE), as shown in the 2010 Digest of Education Statistics (Table 254). Both overall and faculty staffing levels have remained roughly constant during the past three decades. The most notable change has been a reduction in non-professional staff in favor of non-faculty professionals. Assuming that worker quality has stayed constant as well, these constant staffing levels are consistent with our hypothesis that higher education has enjoyed only modest efficiency gains.

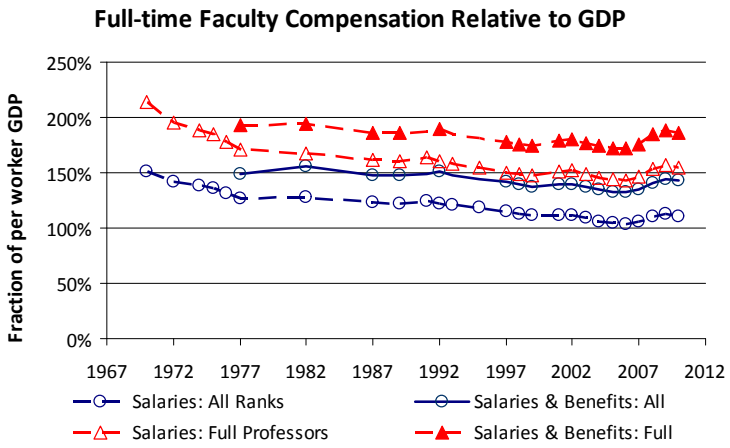

Figure 5

Table 1. Employees per Student			
	1976	1999	2009
All	0.185	0.209	0.186
Professional staff	0.102	0.134	0.132
Administrative	0.012	0.014	0.015
Faculty	0.060	0.067	0.062
Graduate assistants	0.010	0.009	0.009
Other professionals	0.020	0.043	0.046
Non-professional staff	0.084	0.075	0.054
Note: All quantities in FTE terms.			

The Digest of Education Statistics also contains data on faculty compensation. Figure 5 shows that salaries for full-time instructional faculty have fallen slightly relative to GDP. Controlling for faculty rank (full, tenured professors vs. all ranks) and including benefits (not rank-differentiated) does not change the trend. In addition, the fraction of instructional faculty that are full-time employees has fallen, from 78% in 1970 to 51% in 2007. While these trends have almost surely reduced the growth in college costs, interpreting them is difficult. The trends may well reflect a decline in the human capital embodied in college instructors. Alternatively, they may reflect idiosyncratic features in the specialized academic job market. Moreover, it is not clear whether reductions in human capital imply reductions in educational quality or reductions in cost through higher efficiency or the substitution of capital for labor. In the model below, one of our identifying assumptions will be that the quality of higher education is constant.

On the other hand, the reduction of non-professional staff in favor of non-faculty professionals has probably raised staffing costs. Unfortunately, the data do not provide compensation information for non-instructional employees.

2.3 Capital

The Digest of Education Statistics reports total physical plant in the higher education sector for selected years between 1899-1900 through 1994-95. Unfortunately, these data do not divide capital between "education and general uses", as opposed to auxiliary "current fund" uses such as dormitories.[6] Assuming capital is proportional to operating cost, and applying the GDP deflator, Table 2 shows real, educational and general-related, physical plant per FTE, both for all colleges and for the "4-year" institutions we target in our estimation. Table 2 has two notable features. The first is that even though real wages have risen dramatically relative to the cost of capital goods, total capital usage has at most risen modestly over the sample period. This suggests that colleges have at most a limited ability to substitute between capital and labor inputs. The second notable feature is that while the quantity of structures has been roughly constant, the use of equipment has risen significantly. This is consistent with the aggregate capital data, which show that while the relative (to GDP) price of structures has in fact risen slightly, the relative price of equipment has fallen. (See e.g., Cummins and Violante, 2002.) The data thus suggest that capital use will increase over the long run, driven by an increase in equipment.

Table 2. Education & General Capital

	All Colleges	"4-year" Institutions		
	Total Capital	Equipment	Structures	Total Capital
1959-60	31.08	4.18	29.08	33.26
1969-70	30.74	4.39	30.32	34.71
1975-76	31.48	4.90	32.56	37.47
1980-81	32.43	5.38	33.21	38.59
1985-86	31.11	6.28	30.47	36.76
1990-91	30.88	7.27	29.80	37.07

Note: All quantities in 1,000s of $2005 per FTE.

[6] Another limitation of the Department of Education's capital data is that its capital is measured at book, or historical cost. As described in the data Appendix, we convert the book values into current cost values using the Bureau of Economic Analysis' fixed asset data.

2.4 College Attainment and Earnings

Moving to the demand side of the higher education market, Figure 6 presents college attainment by year of birth, using the IPUMS implementation of the Current Population Survey (CPS).[7] Because most people have completed their undergraduate studies by the time they reach 30, we measure each cohort's attainment as the average across ages 28-32. (We measure graduate degree attainment over ages 33-37.) Figure 6 shows that college attainment rose rapidly among the cohorts born between 1935 and 1950; this in part reflects enrollment to avoid the military draft (Card and Lemieux, 2001), and the financial aid provided to Vietnam war veterans (Angrist and Chen, 2011). Attainment fell slightly among the cohorts born between 1950 and 1960, rose rapidly among the cohorts born between 1960 and 1970, and has risen slowly ever since.

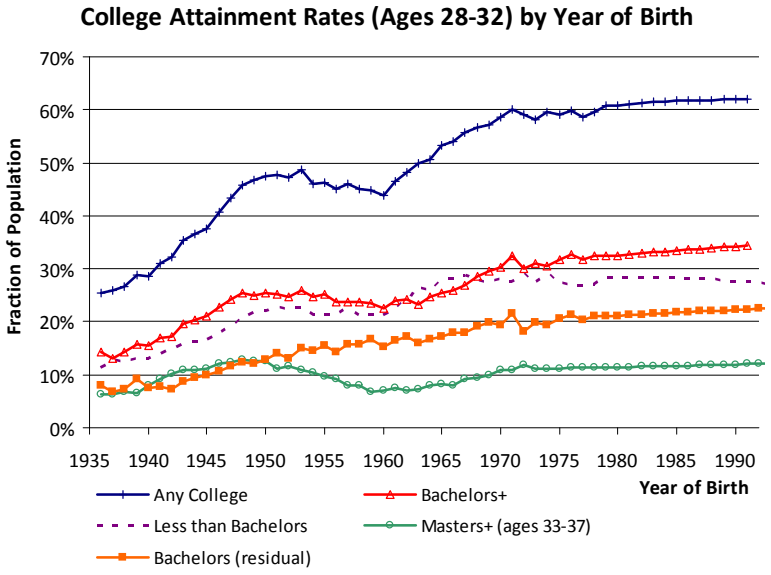

Figure 6

Another piece of the puzzle is earnings. Figure 7 shows the college premium for earnings and wages, again calculated from IPUMS CPS data. We calculate the premium as ratio of the mean earnings (wages) of people aged 30-74 with at least a Bachelor's degree to the mean earnings (wages) of people aged 30-74 whose highest degree is a high school diploma. Figure 7 shows that in 1961 (the March 1962 survey), men with a Bachelor's degree earned 118 percent more than high school graduates. By 2010, the premium had

[7] Assuming the attainment rates have a quadratic trend and an autoregressive residual, we use the data for 1936-1979 to project attainment for people born between 1980 and 1990.

risen to 163 percent. During the same period, the college wage premium, which does not account for hours of work, rose from 58 percent to 93 percent. This rise in college premia is well-known. While the premia reflect differences in ability as well as human capital accumulation, they suggest that the returns to college are probably increasing.

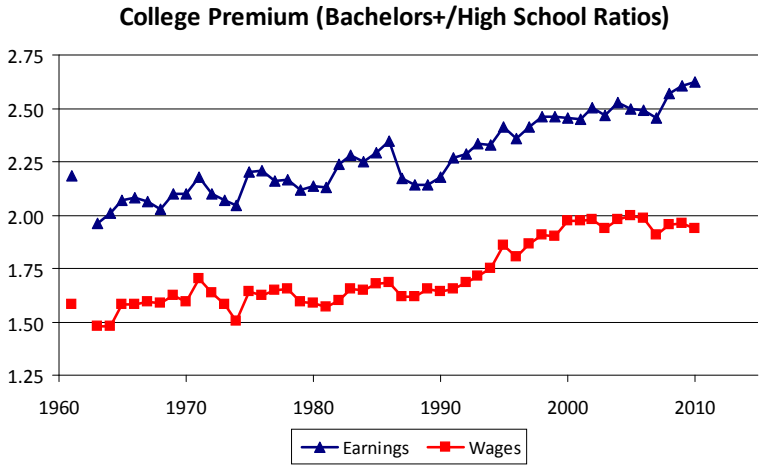

Figure 7

2.5 Key Findings

Our review of the data reveals several trends:

1. Expenditures per college student tend to grow at the same rate as per capita GDP.

2. Since World War II, sticker price tuition has grown faster than GDP, while tuition net of grant aid has grown at the same rate.

3. Since at least 1976, staffing at institutions of higher education has remained constant, although there have been changes in its composition.

4. Capital per student is at most growing slowly over time. Any such growth is being driven by increased use of equipment.

5. Since the 1935 birth cohort, the educational attainment of 30-year-olds, whether measured by attendance or 4-year degrees, has more than doubled.

6. Since 1960, the earnings premium for college-educated workers has grown significantly.

To our knowledge, the first two findings are relatively novel;[8] the third finding is publicly available but not well-known; the fourth finding is novel; and the last two findings are well-known. In the remainder of the paper, we use these findings to motivate and estimate our structural model.

3 Model

We construct a non-stationary OLG model with two sectors: a goods-producing sector and an education sector. The goods sector experiences skill-biased and skill-neutral technical progress, while the education sector does not. The economy also experiences demographic changes. Forward-looking individuals live from ages 18 to 74, and differ by ability. At the beginning of their lives, they decide whether to spend 4 years attending college. The cost of higher education is endogenously determined.

3.1 Goods sector

The goods sector produces output (Y_t) using two skill categories of workers, white-collar and blue-collar, and homogeneous capital. Specifically, production at time t is given by the nested CES function

$$\begin{aligned} Y_t &= A_t K_t^{\alpha} [\omega_t W_t^{1-\varsigma} + (1-\omega_t) B_t^{1-\varsigma}]^{(1-\alpha)/(1-\varsigma)} && \text{(1a)} \\ &= A_t K_t^{\alpha} L_t^{1-\alpha}, && \text{(1b)} \\ L_t &\equiv [\omega_t W_t^{1-\varsigma} + (1-\omega_t) B_t^{1-\varsigma}]^{1/(1-\varsigma)}, && \text{(1c)} \end{aligned}$$

where B_t denotes total units of blue-collar skill, W_t denotes white-collar skill, L_t denotes total labor inputs, K_t denotes capital, and A_t indexes aggregate productivity.[9] The parameter ς governs the substitutability between white- and blue-collar labor (the elasticity of substitution is $1/\varsigma$). To capture skill-biased and skill-neutral technical change, we allow the weight ω_t and the shifter A_t to vary over time.

[8] In an earlier drafts, Castro and Coen-Pirani (2012) also construct an expenditures-to-income measure.

[9] Heckman et al. (1998a) estimate a version of this model where K_t and L_t are also nested in a CES aggregator, but find the elasticity of substitution between K_t and L_t to be close to 1.

Firms are perfectly competitive. The equilibrium pricing conditions are:

$$r_t + \delta = \alpha \frac{Y_t}{K_t} = \alpha A_t \left(\frac{K_t}{L_t} \right)^{\alpha - 1}, \tag{2}$$

$$w_t^B = (1-\alpha)(1-\omega_t) \frac{Y_t}{L_t} \left(\frac{B_t}{L_t} \right)^{-\varsigma}, \tag{3}$$

$$w_t^W = (1-\alpha)\omega_t \frac{Y_t}{L_t} \left(\frac{W_t}{L_t} \right)^{-\varsigma}, \tag{4}$$

where r is the real interest rate, δ is the depreciation rate, and w_t^B and w_t^W are the unit prices of blue-collar and white-collar skill, respectively.

3.2 Individuals

Each year t, a new cohort b $(= t)$ starts its economic life. For $t \geq b$, members of cohort b are of age $a = t - b$; let $N_{a,t}$ denote the number of people of age a at time t. Individuals live from ages 18 $(a = 0)$ to 74 $(a = 56)$, for a total of $a_{\max} = 57$ periods.

3.2.1 Preferences

Each individual's utility depends on two objects. The first is his consumption stream, $\{x_{a,b+a}\}_a$, which he values according to

$$\sum_{a=0}^{a_{\max}-1} \beta^a \frac{1}{1-\varphi} x_{a,b+a}^{1-\varphi},$$

where $0 < \beta < 1$ is the discount factor, and φ is the (inverse) coefficient of relative risk aversion.

The second is the risk of being inducted through the military draft. Throughout most of the Vietnam war, young men could receive a draft deferment by attending college. Card and Lemieux (2001) argue that the fraction of men enrolled at ages 20 to 21 rose and fell in parallel with the risk of being drafted. Figure 6 is consistent with their finding. To account for this effect, we assume that blue-collar workers incur the psychic cost

$$\Delta(i_b) = \Delta_1 i_b + \Delta_2 i_b^2,$$

where i_b denotes the probability of induction for a man in cohort b. The data in Card and Lemieux (2000, Figure 4) show that induction risk is highest for the cohort born in

1946, and vanishes with the end of the draft in the early 1970s.

3.2.2 Earnings and expenditures

Each individual begins life with an idiosyncratic draw of ability, h, from a log-normal distribution: $\ln(h) \sim \mathcal{N}(\mu, \sigma^2)$. Ability varies across individuals, but is constant over their lives. At the beginning of his life, an individual choose whether to spend 4 years attending college. If the individual chooses not to go to college, he works as a unskilled laborer with blue-collar human capital h. If the individual chooses to attend college, after 4 years he becomes a "skilled" worker and his white-collar human capital is

$$\gamma(h) = \gamma_1(h - \gamma_2)^3. \tag{5}$$

The shape of this transformation draws on Heckman et al.'s (1998a, Tables I and II) estimates of how earnings vary by ability and education. Individuals cannot change careers; we are implicitly assuming that unskilled workers have no white-collar human capital, and that skilled workers have no blue-collar human capital.

One of our identifying assumptions is that the distribution of abilities and the transformation function $\gamma(h)$ are both fixed over time. In our model, the earnings of white-collar workers increase because of increases in the price of white-collar skill. An alternative scenario is that the human capital provided by college is increasing over time, so that increases in white-collar earnings are due to increases in the quality of college (Bowlus and Robinson, 2012). In either case, the key technological rigidity is the limited ability of colleges to eliminate white collar workers or replace them with capital. The salaries of higher education employees may rise because their skill has a higher market value, or because they have more skill. Either way, the higher salaries lead to higher college costs. We use our specification because it is easier to interpret and is comparable to much of the existing literature.

The productivity of both blue- and white-collar workers varies exogenously over their life cycles. We allow the profiles themselves to vary over time, to reflect changes in health and labor force participation patterns. Let $\varepsilon^B_{a,t}$ denote the relative efficiency of a blue-collar worker of age a at time t, and let $\varepsilon^W_{a,t}$ denote the corresponding quantity for a white-collar worker. The earnings of a blue-collar individual of type h and age a at time t are

$$y^B_{h,a,t} = h \cdot \varepsilon^B_{a,t} \cdot w^B_t. \tag{6}$$

The earnings of the same individual when she works as white-collar worker are

$$y^W_{h,a,t} = \gamma(h) \cdot \varepsilon^W_{a,t} \cdot w^W_t. \tag{7}$$

The lifetime earnings of a blue-collar worker of cohort b and ability level h can thus be written as $h \cdot J^B_b$, where

$$\begin{aligned} J^B_b &= \sum_{a=0}^{a_{\max}-1} q(b,a)(1-\tau)\varepsilon^B_{a,b+a} w^B_{b+a}, \\ q(b,a) &\equiv \begin{cases} \prod_{j=1}^{a} \frac{1}{1+r_{b+j}(1-\tau)}, & a > 0, \\ 1, & a = 0 \end{cases}, \end{aligned}$$

and τ is the income tax rate. The lifetime earnings of a white-collar worker are the sum of two elements: $\chi_0 h \cdot J^{WB}_b$, income earned during college as a blue-collar worker; and $\gamma(h) \cdot J^{WW}_b$, income earned as a white-collar worker after graduation:

$$\begin{aligned} J^{WB}_b &= \sum_{a=0}^{3} q(b,a)(1-\tau)\varepsilon^B_{a,b+a} w^B_{b+a}, \\ J^{WW}_b &= \sum_{a=4}^{a_{\max}-1} q(b,a)(1-\tau)\varepsilon^W_{a,b+a} w^W_{b+a}. \end{aligned}$$

The parameter $0 \leq \chi_0 \leq 1$ reflects the student's employment load during college.

The cost of a year of college at time t is c_t . Students pay the fraction d_t of this cost, with the remainder funded out of tax revenues. Let $C_b = \sum_{a=0}^{3} q(b,a) d_{b+a} c_{b+a}$ denote the lifetime cost of college for cohort b. We assume that college students can borrow only a fraction of their future earnings, and estimate the fraction using our model. In particular, we assume that

$$C_b + \sum_{a=0}^{3} q(b,a) x^W_{h,a,b+a} - \chi_0 h J^{WB}_b \leq \chi_1 \gamma(h) J^{WW}_b, \tag{8}$$

where $0 \leq \chi_1 \leq 1$ is the fraction of post-graduation earnings available for borrowing. This lending restriction is similar to the constraint for private lending found in Lochner and Monge-Naranjo (2011, equation (7)), although it should be interpreted more broadly. The limit increases with the individual's skill level. Lochner and Monge-Naranjo (2011) show that borrowing limits that do not vary with future earnings can imply that high-

ability people are *less* likely to attend college, the opposite of the empirical relationship.

To simplify the model, we assume that college attendees can smooth consumption perfectly during their 4 years of college, and during their working lives. Equation (8) restricts their ability to smooth consumption between the two phases of their lives, college and work, but not within each phase. College attendees thus maximize their lifetime utility subject to equation (8) and the lifetime budget constraint

$$\sum_{a=0}^{a_{\max}-1} q(b,a)x^W_{h,a,b+a} \leq \chi_0 h J^{WB}_b + \gamma(h) J^{WW}_b - C_b. \tag{9}$$

Blue-collar workers solve a similar problem, but with no education expenditures they are able to smooth perfectly over their entire lives: their only constraint is an analogue of equation (9). Recall that h is a lifetime constant and there are no aggregate shocks. With perfect foresight and homothetic preferences, the consumption allocations for both types of workers, and the resulting lifetime utilities, can be expressed analytically.

3.2.3 The college decision

In deciding whether to attend college, new workers compare potential consumption streams and psychic costs. This leads to a threshold ability level $h^*_b \geq \hat{h}_b$ such that an individual chooses to go to college if and only if $h > h^*_b$.[10] Let

$$e_b = 1 - \Phi\left(\frac{\ln(h^*_b) - \mu}{\sigma}\right) \equiv 1 - F(h^*_b), \tag{10}$$

where $\Phi(\cdot)$ is the standard normal distribution function, denote the fraction of cohort b that attends or has attended college.

In our model college enrollment and attainment are identical; our model does not account for drop-out risk. Our distribution of skill thus captures both differences in "ability" and differences in drop-out risk. Bound, Loevenheim, and Turner (2009) show that college completion rates drop markedly as one moves down the distribution of AFQT scores.

[10] The existence of a single threshold ability level is not guaranteed. However, checks of our numerical solutions indicate that in practice the enrollment decision is characterized by a single threshold.

3.3 Higher education sector

Converting a blue-collar worker into a white-collar worker requires skilled labor and capital. Because markets are competitive, each of these inputs must be paid their market price, namely the wage w_t^W and the user cost $(r_t + \delta)$. The cost of educating a college student for one year is

$$\begin{aligned} &\min \; W_t^E w_t^W + K_t^E (r_t + \delta), \\ &s.t. \; 1 = z(\kappa (W_t^E)^{1-\upsilon} + (1-\kappa)(K_t^E)^{1-\upsilon})^{1/(1-\upsilon)}, \end{aligned} \tag{11}$$

where W_t^E denotes white-collar skill employed in higher education and K_t^E denotes higher education capital. We have normalized the "quantity" of education to 1, so that the productivity factor z scales education costs. The parameter υ governs the substitutability between capital and white-collar labor (the elasticity of substitution is $1/\upsilon$).

We assume that colleges seek to minimize costs, and operate on their production frontier. We do not doubt that the higher education sector has inefficiencies. However, our data show that between 1959 and 2009, real expenditures per student for our "4-year" institutions rose 163%. If inefficiency is the main culprit, colleges must have gotten much *more* inefficient over time. An increase in inefficiency of this scale is possible only if institutions of higher education do not actively compete. But colleges compete avidly to achieve higher U.S. News and World Report rankings (Monks and Ehrenberg, 1999). Moreover the supply of higher education does respond to increased demand. The number of college students has risen dramatically over time, both in absolute terms and as at fraction of the population.

Returning to the model, the first order conditions for cost minimization are

$$W_t^E = \left(\frac{\kappa c_t}{w_t^W} \right)^{1/\nu} z^{(1-\nu)/\nu}, \tag{12}$$

$$K_t^E = \left(\frac{(1-\kappa) c_t}{r_t + \delta} \right)^{1/\nu} z^{(1-\nu)/\nu}, \tag{13}$$

where c_t is the multiplier on equation (11) and (equivalently) the total cost of college. Substituting these results into equation (11) yields

$$c_t = z^{-1} \left[\kappa^{1/\nu} (w_t^W)^{(\nu-1)/\nu} + (1-\kappa)^{1/\nu} (r_t + \delta)^{(\nu-1)/\nu} \right]^{\nu/(\nu-1)}. \tag{14}$$

It is useful to consider what happens in the limit as w_t^W becomes arbitrarily large.

For $\nu < 1$, the Inada conditions do not hold. Labor inputs go to zero and equation (14) simplifies to

$$c_t = z^{-1}(1-\kappa)^{1/(\nu-1)}(r_t+\delta).$$

For $\nu > 1$, equations (14) and (12) simplify to

$$\begin{aligned} c_t &= z^{-1}\kappa^{1/(\nu-1)} w_t^W, \\ W_t^E &= z^{-1}\kappa^{1/(\nu-1)}. \end{aligned}$$

3.4 Equilibrium

We will work with an open economy framework, taking the sequence of interest rates $\{r_t\}$ as given. We assume that any tax revenues not spent on education subsidies are used to purchase government goods, which affect neither production nor utility. Recall that $N_{a,t}$ denotes the number of people of age a at time t.

Definition. An equilibrium consists of sequences of: capital and labor inputs $\{K_t, W_t, B_t, K_t^E, W_t^E\}$, wage rates $\{w_t^B, w_t^W\}$, costs $\{c_t\}$, and skill thresholds $\{h_t^*\}$ such that:

(i) Given $\{r_t, w_t^B, w_t^W, c_t\}$, $\{h_t^*\}$ solves the individual's problem.

(ii) The price of each factor equals its marginal product. That is, equations (2), (3) and (4) hold.

(iii) Colleges earn no profits: equation (14) holds.

(iv) All markets clear:

$$W_t = \sum_{a>3} N_{a,t}\,\varepsilon_{a,t}^W \int_{h>h_{t-a}^*} \gamma(h)\,dF(h) - W_t^E \sum_{a\leq 3} N_{a,t}\,e_{t-a}, \tag{15}$$

$$B_t = \sum_{a} N_{a,t}\,\varepsilon_{a,t}^B \int_{h\leq h_{t-a}^*} h\,dF(h) + \chi_0 \sum_{a\leq 3} N_{a,t}\,\varepsilon_{a,t}^B \int_{h>h_{t-a}^*} h\,dF(h), \tag{16}$$

4 Estimation

We solve the model for the period 1961-2300. We assume that TFP (A_t) grows at a constant rate through the entire period, and that after 2200 the population $\{N_{t-a,t}\}_a$ is constant, so that the economy converges to a balanced growth path.[11] We find the

[11] Because it incorporates the cost of educating a college student, the balanced growth path is asymptotic. As described in section 3.3 above, in the limit the cost of college either: converges to a constant that becomes irrelevant as the economy continues to grow; or converges to the white-collar wage, which follows the balanced growth path.

equilibrium path with a guess of the sequence of blue- and white-collar wages $\{w_t^B, w_t^W\}$. Using this guess, we find educational attainment for each cohort, which in turn allows us to find the aggregate supplies of blue- and white-collar skill for each year. These skill totals in turn provide us with new estimates of wages, calculated as marginal products. We search over wage sequences until we reach a fixed point.

Following a number of papers (e.g., French, 2005; De Nardi, French and Jones, 2010), we find the parameters of the model using a two-step process. First, we calibrate or estimate a number of processes outside of the model. For example, we calibrate the demographic transition using Census projections. Additional details can be found in Appendix A. Second, we estimate the remaining parameters with a variant of non-linear least squares, taking the first-stage parameters as given. Because the targets we match in this second step can be viewed as time series realizations, the asymptotic properties of our estimates follow from standard Method of Moments arguments. Additional details can be found in Appendix B.

4.1 Processes and Parameters Estimated Outside the Model

Table 3 provides a summary of how the first-step parameters are set.

4.1.1 Demographics

We begin with Census data (inclusive of Armed Forces overseas) for the period 1961-2000. For the period 2000-2050, we use the Census Bureau's 2008 projections. These detailed projections allow us to calculate "survival" rates (net of immigration) at every age and the growth rate of the birth cohorts. We project all of these rates for the period 2051-2100 by linearly extrapolating their values over the period 2040-2050, and use the projected rates to update the population. From 2100 forward we assume that the size of the birth cohort is constant,[12] and that the net survival rates are constant as well. By 2200 the population has converged to a stationary distribution. Our calculations use the subset of the age distribution covering ages 18-74.

[12] In this respect, we follow Krueger and Ludwig (2007).

Table 3. Parameters Found without Using the Model

Parameters		Values/Source
$\{N_{a,t}\}$	Demographics	Census Bureau (section 4.1.1)
$\{\varepsilon^B_{a,t}, \varepsilon^W_{a,t}\}$	Life-cycle efficiency profiles	CPS (section 4.1.2)
$\{d_t\}$	Net tuition copay rate	Dept. of Education (section 4.1.3)
$\{e_b\}_{b<1943}$	Education, older cohorts	CPS (section 4.1.4)
$\{i_b\}$	Military induction risk	Card and Lemieux (2000)
δ	Capital depreciation rate	0.05, standard
β	Discount factor	0.97, standard
α	Capital share in production	0.3, standard
φ	1/(coefficient of RRA)	2, standard
τ	Income tax rate	$0.2 \approx$ average G/Y (section 4.1.5)
ς	1/(skill substitutability)	0.7, Heckman et al. (1998a)
χ_0	Fraction of time students spend working	0.5 (section 4.1.5)
$\{r_t\}$	real interest rates	National accounts, using equation (2)
g_A	TFP growth rate	0.75%, Jorgenson and Yip (2001)
A_{1960}	Logged initial TFP level	1, normalization (section 4.1.5)

4.1.2 Efficiency Profiles

Taking averages over equations (6) and (7), we can find mean blue-collar and white-collar earnings for each age-year-cohort combination:

$$\begin{aligned} \overline{y}^B_{a,t} &= E(h|h \leq h^*_b) \cdot \varepsilon^B_{a,t} \cdot w^B_t, \\ \overline{y}^W_{a,t} &= E(\gamma(h)|h > h^*_b) \cdot \varepsilon^W_{a,t} \cdot w^W_t. \end{aligned}$$

where $b = t - a$ indexes the agent's birth cohort, and h^*_b is the education cut-off level for cohort b. The model will generate the skill levels $E(h|h \leq h^*_b)$ and $E(\gamma(h)|h > h^*_b)$ and the skill prices w^B_t and w^W_t. The life-cycle efficiency profiles $\{\varepsilon^B_{a,t}, \varepsilon^W_{a,t}\}$, on the other hand, are treated as exogenous, and need to be estimated from the data.

Focussing first on blue-collar workers, we have CPS data for blue-collar (high school graduate) earnings. Taking averages gives us $\{\widetilde{y}^B_{a,t}\}$, a synthetic panel (Deaton, 1985). Because we are interested in overall earnings, including those of non-workers, the averages

are taken across all individuals. Continuing, we decompose $\ln(\widetilde{y}^B_{a,t})$ as

$$\begin{aligned}
\ln(\widetilde{y}^B_{a,t}) &= \ln(\overline{y}^B_{a,t}) + \eta^B_{a,t}, \\
&= \ln\left[E(h|h \leq h^*_b)\right] + \ln(\varepsilon^B_{a,t}) + \ln(w^B_t) + \eta^B_{a,t}, \\
\ln\left[E(h|h \leq h^*_b)\right] &= \zeta^B_1 e_b + \zeta^B_2 e_b^2 + \zeta^B_3 e_b^3 + \zeta^B_4 e_b^4, \\
\ln(w^B_t) &= \sum\nolimits_u \lambda^B_u \cdot 1\{t = u\}, \\
\ln(\varepsilon^B_{a,t}) &= \xi^B_0 + \xi^B_1 a + \xi^B_2 a^2 + \xi^B_3 a^3 + \xi^B_4 t \cdot a + \xi^B_5 t \cdot a^2 + \xi^B_6 t_2 \cdot a + \xi^B_7 t_2 \cdot a^2,
\end{aligned}$$

where: $\eta^B_{a,t}$ is classical measurement error; e_b denotes the fraction of cohort b with a 4-year college degree; $1\{t = u\}$ is a time dummy; and $t_2 = \max\{t - 1990, 0\}$ is a spline term. Since the ability threshold h^*_b determines the attainment level e_b through equation (10), we can approximate $\ln\left[E(h|h \leq h^*_b)\right]$ arbitrarily well with a polynomial in e_b. We thus handle the standard age-time-cohort problem by assuming, à la Carneiro and Lee (2011), that cohort effects can be expressed as a function of educational attainment. The changes in skill prices are assigned to the time dummies. This leaves the age-efficiency profile.

Following Kambourov and Manovskii (2009), we assume that the age-efficiency relationship shifts over time, using the coefficients $\{\xi^B_4, \xi^B_5, \xi^B_6, \xi^B_7\}$. Because we are working with earnings data, $\ln(\varepsilon^B_{a,t})$ will shift with changes in labor supply. Two sources of shifts are changes in retirement (French and Jones, 2012) and increased participation by women. The earnings profiles might also vary because of changes in on-the-job human capital accumulation (Heckman et al., 1998a). All these mechanisms are outside our model, and thus included in the age-efficiency profiles.

Collecting the preceding equations, we thus estimate

$$\begin{aligned}
\ln(\widetilde{y}^B_{a,t}) &= \xi^B_0 + \xi^B_1 a + \xi^B_2 a^2 + \xi^B_3 a^3 + \xi^B_4 t \cdot a + \xi^B_5 t \cdot a^2 + \xi^B_6 t_2 \cdot a + \xi^B_7 t_2 \cdot a^2 \\
&\quad + \zeta^B_1 e_b + \zeta^B_2 e_b^2 + \zeta^B_3 e_b^3 + \zeta^B_4 e_b^4 + \sum\nolimits_u \lambda^B_u \cdot 1\{t = u\} + \eta^B_{a,t}, \qquad (17)
\end{aligned}$$

and use the right-hand-side terms on the first line as our estimate of $\{\varepsilon^B_{a,t}\}$. The earnings profiles for white collar workers are found in a similar fashion.

Figure 8 shows the efficiency profiles for workers born in 1917, 1942 and 1966. The underlying coefficients can be found in Appendix A. It bears noting that the absolute levels of the efficiency profiles (ξ^B_0, ξ^W_0) are not identified separately from the parameters of the skill distribution: it is the products $h \cdot \varepsilon^B_{a,t}$ and $\gamma(h) \cdot \varepsilon^W_{a,t}$ that determine earnings potential. In Figure 8 we have chosen to normalize the levels of the profiles to those of observed earnings. However, it is only the life-cycle variation within these profiles that

affects our results.

Like Kambourov and Manovskii (2009), we find that earnings profiles flatten over time. The changes are most notable after age 60. We assume that efficiency profiles continue to evolve in this fashion until the 1970 cohort is reached, after which they remain constant.

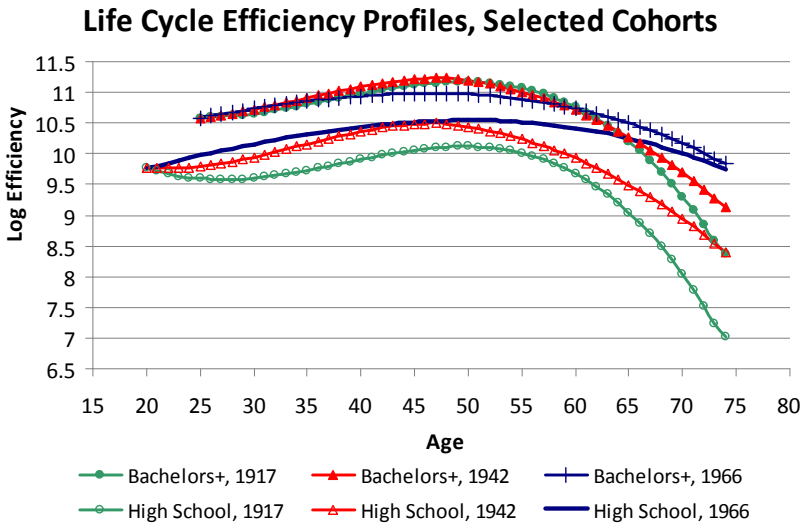

Figure 8

These profiles aggregate male and female earnings. In doing this, we are assuming that once one controls for age-efficiency profiles, men and women make education choices in the same fashion; aggregating across genders is no different than any other form of aggregation. This is consistent with Ge and Yang (2013), who find that most of the increase in female college attainment can be explained by increased financial returns, reflecting in part higher female labor market participation. The underlying changes in social norms and home production technologies that increase female earnings are picked up in the age-efficiency profiles.

4.1.3 Financial Aid

Our measure of the tuition co-pay rate, d, is the ratio of net tuition to total costs, both of which are taken from the data described in Section 2. To be consistent with our modelling of the education decision, we base our measure on data for "4-year" institutions. Since 1985, d has ranged between 20 and 25%. We assume that d remains at its 2009 value of 23.2% through 2300. Missing values for earlier years are linearly interpolated. Additional details can be found in Appendix A.

4.1.4 Educational Attainment, Older Cohorts

The educational attainment, the fraction of the population with at least a Bachelors degree, of people 19 and older in 1961 – cohorts born before 1943 – is taken from the CPS. When possible, we measure education attainment as the average over ages 28-32; for older cohorts we take averages across the 5 (if available) youngest ages observed in our data.

4.1.5 Miscellaneous Parameters

We set the depreciation rate $\delta = 0.05$; the aggregate capital share $\alpha = 0.3$; the discount factor $\beta = 0.97$; the utility curvature parameter $\varphi = 2$; the tax rate $\tau = 0.2$; and the labor substitutability parameter $\varsigma = 0.7$. The first three values are standard; the fourth is well within the range of micro-level estimates; the fifth is (roughly) the share of government goods and services in GDP since 1960; and the sixth is taken from Heckman et al. (1998a, Table III, OLS). We set χ_0, the fraction of time a student spends working, to 0.5. This reflects the notion that students work full time during the three summer months, and 1/3 time during the rest of the year. Our measure of induction risk, i_b, comes from Card and Lemieux (2000, Figure 4).[13] We construct the interest rate series $\{r_t\}$ by solving equation (2) with the nominal output/capital ratio.[14]

We assume that the log of TFP follows a linear trend: $\ln(A_t) = A_{1960} + g_A t$. We set g_A to 0.75%, the quality-adjusted rate found by Jorgenson and Yip (2001, Table 12.6, 1960-95). As Lee and Wolpin (2006, 2010) point out, A_{1960} is not identified independently of the average level of the skill totals W_t and B_t. We normalize A_{1960} to 1, and estimate the skill distributions behind W_t and B_t within the model.

4.2 Parameters Estimated within the Model

4.2.1 Econometric Strategy

The following parameters are estimated within the model:

1. The skill weights, $\{\omega_t\}$. Building on Heckman et al., (1998a), we assume that ω_t

[13] We are grateful to Thomas Lemieux for providing the data behind this figure.

[14] Let $q_t = p_{Kt}/p_{Yt}$ denote the relative price of capital. If q_t is allowed to differ from 1, the first-order condition for capital is $q_t[1 + r_t - \frac{q_{t+1}}{q_t}(1-\delta)] = \alpha Y_t/K_t$. Assuming further that $\frac{q_{t+1}}{q_t} \approx 1$, this simplifies to $r_t + \delta = q_t^{-1}\alpha Y_t/K_t = \alpha(p_{Yt}Y_t)/(p_{Kt}K_t)$, which uses the nominal output/capital ratio.

follows a logistic trend:

$$\omega_t = \iota_3 + \frac{\exp(\iota_0 + \iota_1 t)}{1 + \exp(\iota_0 + \iota_1 t)}[\iota_2 - \iota_3], \tag{18}$$

with the upper bound $\iota_2 \in (0, 1)$. We estimate ι_0 through ι_3.

2. The input weight κ and the substitution parameter υ in the higher education production function.

3. TFP in the production of higher education, z.

4. The parameters of the skill transformation function, equation (5), γ_1 and γ_2.

5. The parameters of the lognormal skill distribution, $\ln(h) \sim \mathcal{N}(\mu, \sigma^2)$.

6. The borrowing constraint parameter χ_1.

7. The induction cost parameters Δ_1 and Δ_2.

To find these parameters, we match the following observations:

1. The educational attainment – the fraction of the population with at least a Bachelors degree – of each of the cohorts born between 1943 and 1991.

2. The college earnings premium, as measured in Figure 7 in Section 2, for each of the years 1961-2009; the model analog to this quantity is mer_t, the ratio of average earnings for college-educated workers to the average earnings for blue-collar workers among people aged 30-74:

$$mer_t = \frac{w_t^W \left[\sum_{a>11} N_{a,t}\, \varepsilon_{a,t}^W \int_{h>h_{t-a}^*} \gamma(h)\, dF(h)\right] / \left[\sum_{a>11} N_{a,t} e_{t-a}\right]}{w_t^B \left[\sum_{a>11} N_{a,t}\, \varepsilon_{a,t}^B \int_{h<h_{t-a}^*} h\, dF(h)\right] / \left[\sum_{a>11} N_{a,t}(1 - e_{t-a})\right]}. \tag{19}$$

3. GDP for each of the years 1961-2009.

4. Expenditures per student by "4-year" institutions for each of the years 1961-2009.[15]

5. Capital per student for "4-year" institutions for each of the years 1961-1994.

[15] To be consistent with our data targets, our measure of educational cost excludes interest expenditures.

In most cases, all we have are the aggregate moments for each date, not the disaggregated underlying data used to construct them. For example, we have a time series of GDP, but not the detailed data used by the Bureau of Economic Analysis to construct it. Our econometric strategy is thus similar to that of Lee and Wolpin (2006, 2010) in that our data consist of aggregate moments, rather than micro-level observations. The key insight is that each moment can be considered a time series observation.

Let f_{mt} denote a observation of type m, $m \in \{1, 2, .., M\}$, in year t, $t \in \{1, 2, ..., T\}$. For example, f_{c1969} would be expenditures per student at "4-year" institutions (item 4 immediately above) in 1969. The model-predicted value of f_{mt} is $f^*_{mt}(\theta)$, where θ is the parameter vector we wish to estimate.[16] The model prediction differs from the data because of measurement error and because of events, such as business cycles, that we do not explicitly model. We capture these effects in a residual term. At the "true" parameter vector θ_0 we assume that

$$f_{mt} = f^*_{mt}(\theta_0)(1 + u_{mt})^{-1},$$

where u_{mt} is, conditional on $f^*_{mt}(\theta_0)$, a mean-zero stationary random variable. Our estimate of θ_0, $\widehat{\theta}$, solves

$$\min_{\theta} \sum_{t=1}^{T} \sum_{m=1}^{M} \left(\frac{f^*_{mt}(\theta)}{f_{mt}} - 1 \right)^2 = \min_{\theta} \sum_{t=1}^{T} \sum_{m=1}^{M} u^2_{mt}. \tag{20}$$

In Appendix B, we derive the asymptotic properties of $\widehat{\theta}$ using standard GMM arguments.

4.2.2 Parameter Estimates, Model Fit and Identification

Table 4 shows the parameter estimates, along with standard errors. We show estimates for three cases: (1) the baseline case described above; (2) the model with no borrowing constraint ($\chi_1 = 1$); and (3) the model with no borrowing constraints and no induction costs ($\Delta_1 = \Delta_2 = 0$).

[16] Because we use an explicit formula for evaluating integrals (see Jawitz, 2004), our model predictions are exact. Our estimates therefore do not suffer from the bias that plagues simulated non-linear least squares (see, e.g., Gourieroux and Monfort, 1996.)

Table 4. Parameters Estimated Using the Model

Parameters		Baseline	No Borrowing Constraint	No Induction Cost
γ_1	skill production function, cubic weight	0.0052 (0.0000)	0.0052 (0.0000)	0.1334 (0.0001)
γ_2	cubic shifter in skill production function	7.278 (0.0067)	7.237 (0.0017)	22.329 (0.0513)
μ	mean of log ability	3.399 (0.0069)	3.397 (0.0007)	3.384 (0.0027)
σ	standard deviation of log ability	0.1676 (0.0004)	0.1696 (0.0000)	0.0730 (0.0000)
ι_0	evolution of skill weight ω, intercept	−2.530 (0.0002)	−2.595 (0.0001)	−2.555 (0.0001)
ι_1	evolution of skill weight ω, time trend	0.0960 (0.0004)	0.0999 (0.0001)	0.0973 (0.0000)
ι_2	upper bound on ω	0.4728 (0.0003)	0.4648 (0.0001)	0.4701 (0.0013)
ι_3	shift parameter for ω	0.1535 (0.0003)	0.1557 (0.0001)	0.1554 (0.0001)
υ	substitutability parameter, education production function	3.551 (0.0039)	3.557 (0.0024)	3.665 (0.0013)
κ	input weight, education production	9.78×10^{-8} (1.15×10^{-10})	9.57×10^{-8} (2.35×10^{-10})	4.37×10^{-8} (1.12×10^{-10})
z	education productivity level	−9.969 (0.0077)	−9.970 (0.0081)	−9.992 (0.0123)
χ_1	tightness of borrowing constraint	0.0782 (0.0004)	1 (NA)	1 (NA)
Δ_1	induction cost, linear component	0.0017 (0.0000)	0.0017 (0.0000)	0 (NA)
Δ_2	induction cost, quadratic component	−0.0111 (0.0000)	−0.0113 (0.0000)	0 (NA)
GMM criterion (equation 20)		0.6012	0.6014	0.7090

Figure 9 plots the skill weight ω_t for the baseline model. Starting from a value of 0.15 in 1961, ω_t increases dramatically, before levelling off at around 0.47 by 2100. The other versions of the model have similar trajectories. The path of the skill weight is identified by the trajectories of educational attainment and the mean earnings ratio. As ω_t rises, all else equal, white-collar skill will become more productive, raising white-collar wages, the returns to college, and attainment. The effect of ω_t on the earnings ratio is less clear. Although increases in ω_t raise the relative wages of white-collar workers, the increase in educational attainment decreases the average ability of both blue- and white-collar workers, altering the ratio through selection effects (Carniero and Lee, 2011; Hendricks

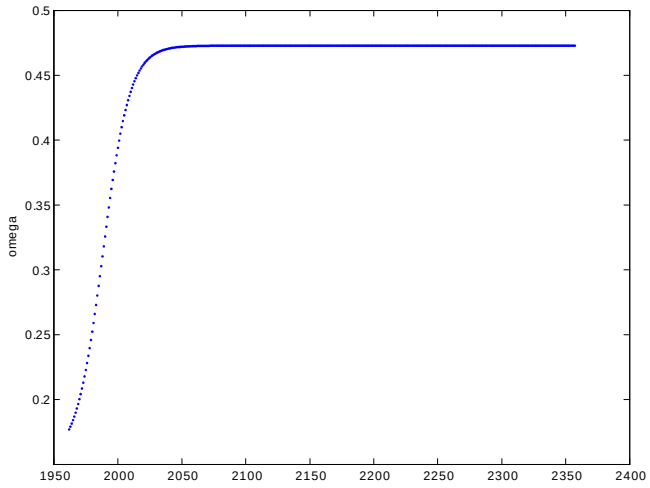

Figure 9: Estimated skill weight

and Schoellmann, 2014).

The same targets also identify the parameters of the ability distribution, $\ln(h) \sim \mathcal{N}(\mu, \sigma^2)$, and the skill transformation function $\gamma(h) = \gamma_1(h - \gamma_2)^3$. Holding attainment fixed, increasing σ^2 imparts an upward skew to the ability distribution, increasing the mean earnings ratio. Similarly, lowering γ_2 increases the white-collar skill acquired by the most able, raising their earnings and increasing the earnings ratio. To sharpen identification, we give the ability distribution and the skill transformation function parsimonious specifications.

The parameter χ_1, which determines the tightness of the borrowing constraint, is identified by time series variation, in particular the degree to which college attainment varies with the college co-pay rate d_t and the college earnings premium. Identification is concentrated in two periods. First, attainment rose rapidly in the late 1960s, peaked in the early 1970s, and then fell. The co-pay rate d_t, however, did not fall until the mid-1970s.[17] This suggests loose constraints. The second key period for identification is the past 20 years, where enrollment rose slowly even as the college earnings premium continued to rise (Goldin and Katz, 2008). This suggests tight constraints. Our baseline estimate of χ_1 implies that college students can borrow around 7.8 percent of their future

[17] This drop in the co-pay rate is driven primarily by military benefits enjoyed by Vietnam veterans. Angrist and Chen (2011) find that these benefits significantly increased enrollment among veterans. This raises the possibility that the Vietnam-era cohorts earned their undergraduate degrees at older ages. However, the CPS data, and the work of Card and Lemieux (2001) provide little evidence that overall undergraduate attainment was delayed during the Vietnam era. In addition, Figure 6 shows a signficant increase in graduate-level attainment among the Vietnam cohorts; our measure of financial aid combines undergraduate and graduate assistance.

incomes.

The cross-sectional evidence on whether borrowing limits reduce access to college is mixed. A number of studies, including Cameron and Heckman (1998, 2001), Keane and Wolpin (2001), Carniero and Heckman (2002), and Cameron and Taber (2004) conclude that borrowing constraints did not significantly restrict college attendance. In contrast, Lochner and Monge-Naranjo (2011) and Abbott et al. (2013) find that although credit constraints might not have restricted access in the past, they probably restrict access now. (Also see Ionescu, 2009.)

The cost parameters Δ_1 and Δ_2 are identified by the comovement between attainment and induction risk. Card and Lemieux (2001) show that the surge in the induction risk during the late 1960s was accompanied by a surge in undergraduate enrollment (and attainment). Including induction costs also allows the model to support a smaller value of χ_1. If the surge in enrollment during the late 1960s can be reconciled with regular tuition growth, a tighter borrowing constraint can fit the data.

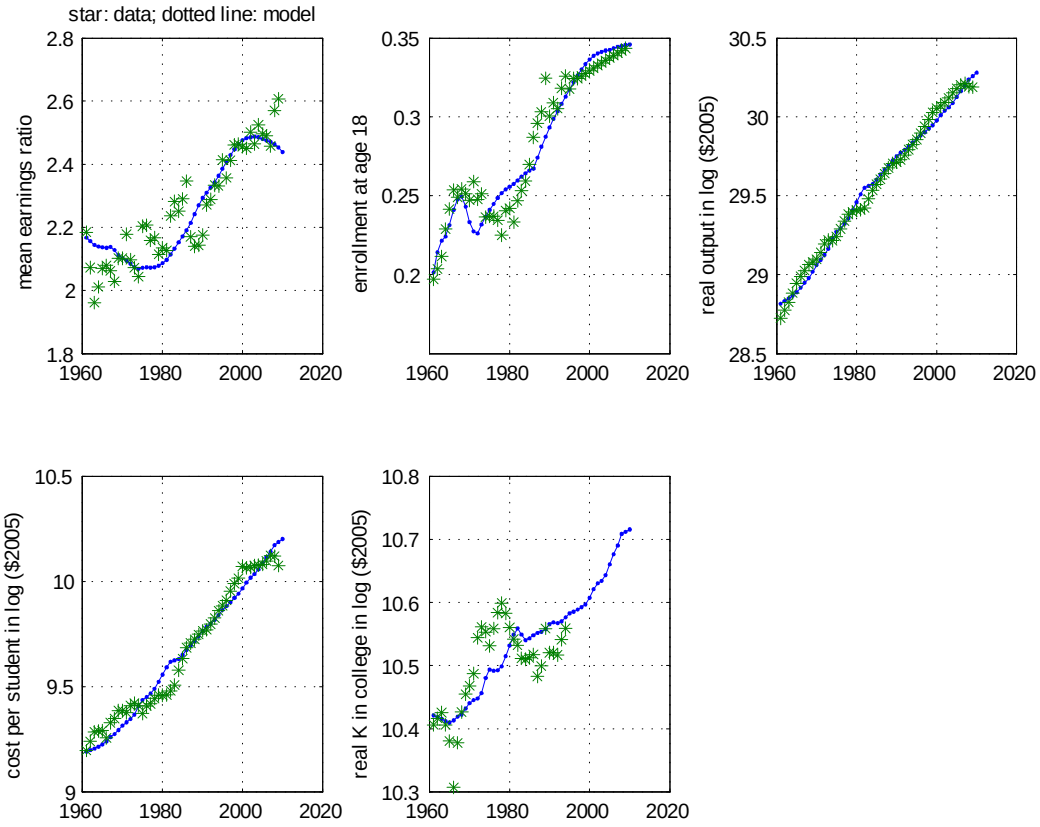

Figure 10: Fit of estimation targets: Baseline model

Figure 10, which shows the model's fit of the estimation targets, illustrates these me-

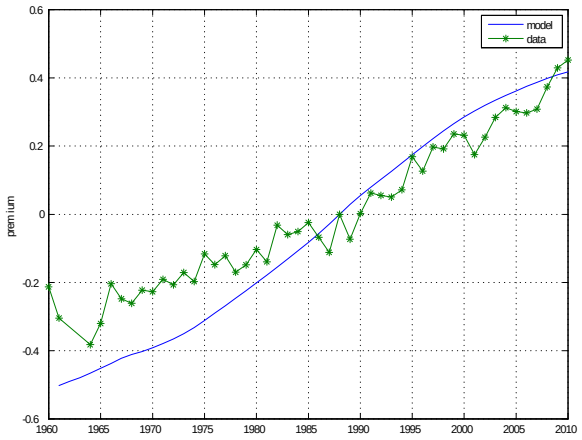

Figure 11: Logged skill premium, data and baseline model

chanics. The two top left panels shows the model's ability to reconcile continued growth in the earnings premium with the slowdown in enrollment. In our framework, matching such a development requires an increase in the demand for skilled labor accompanied by either: (i) negative skill supply shocks; (ii) or a steep skill supply curve. All versions of our model rely heavily on a steep skill supply curve, which they achieve by having the skill level of the marginal college graduate drop rapidly with enrollment, reflecting the cubic skill transformation function.[18] Our models also impose a negative skill supply shock through the increase in net tuition. The strength of this shock, however, depends on the tightness of the borrowing constraint. Holding enrollment growth fixed, the earnings premium can grow more quickly when the borrowing constraint is tight.

The evolution of the mean earnings ratio depends on both the skill premium, i.e. the relative price of skilled labor (w_t^W/w_t^B), and the average abilities of blue- and white-collar workers. Because we explicitly model (and match) both college attainment and the mean earnings ratio, the model-generated earnings premium should reflect the same selection dynamics as the empirical earnings premium. Figure 11 compares the logged skill premia generated by the data – the time dummies in equation (17) – to those generated by the model. Although the skill premia are not estimation targets, the two profiles are similar, giving us confidence in our estimated skill distribution. Both the data and the model show that the skill premium rises more quickly than the mean earnings ratio, suggesting that composition effects reduce the earnings ratio.

[18] A rapid decline in skill is also consistent with the evidence on drop-out risk, which is otherwise not captured in our model. Athreya and Eberly (2013) show that even if all college graduates have similar earnings, empirically plausible changes in drop-out risk will lead to large changes in the expected returns to college.

Figure 10 also shows that the induction cost significantly improves the model's fit of the enrollment profile. The induction costs generate a jump in enrollment during the late 1960s, which in turn depresses the skill premium, leading to slightly lower enrollment in the 1970s.

The parameter υ, which determines the elasticity of substitution in the higher education production function, is identified by: (1) the degree to which capital inputs increase as white-collar skill becomes more expensive, via equations (13) and (14); (2) the degree to which higher education costs grow more slowly than the price of skilled labor. Given that college capital use has grown only modestly over the sample period, the estimated value of 3.55, which suggests that capital and labor are not close substitutes, is not surprising. Even this small degree of substitutability, however, allows colleges to offset some of the effect of higher white-collar wages. This may be reflected in the data, which show slow growth in faculty salaries and increasing use of part-time instruction, along with growth in educational sector equipment. The weighting parameter κ and the technology scaler z are identified by average expenditures and capital inputs for "4-year" institutions. We sharpen identification by holding these parameters constant across time. All three versions of the model do a good job of fitting the cost and capital profiles for the higher education sector. The service sector disease is more than capable of explaining the growth in higher education expenditures.

5 Experiments

In this section we use the model to perform three sets of experiments. In the first, we decompose the changes in the economy between 1961 and 2010 by holding individual parameters fixed at their 1961 values. In the second set of experiments, we determine how college enrollment would have evolved had higher education costs remained at their 1961 levels. In the third set of experiments, we measure how temporary and permanent changes to the tuition co-pay rate would have affected the economy, and assess the role of endogenous costs in generating these effects.

5.1 Decomposition

From 1961 onward, six sets of parameters changed: the skill weight ω, the productivity parameter A, the tuition discount factor d, the lifecycle earnings profiles $\{\varepsilon^W_{a,}, \varepsilon^B_a\}$, the population distribution $\{N_a\}$, and interest rate $\{r_t\}$. To assess the effects of the parameters, we perform a decomposition exercise. The results are summarized in Table 5.

The first two rows of Table 5 compare the economy at 1961 and 2010 under the baseline model. With a higher level of TFP, output is much higher in 2010. College enrollment is also much higher, due to a higher skill premium (w^W/w^B), reflecting a larger weight on white-collar skill in the aggregate production function. The higher skill premium also leads to relatively higher college expenditures: c/y increases from 0.338 to 0.415.

In the final five rows of Table 5, we fix the listed sets of parameters at their 1961 values, allowing the other parameters to evolve as in the baseline model, and evaluate the economy at 2010. Comparing these alternative scenarios to the full model identifies the effects of each parameter shift.

The decomposition shows that the increase in enrollment is due almost entirely to skill-biased technical change. When the skill weight ω increases from its 1961 value, the marginal product of blue-collar workers falls and that of white-collar workers rises. This significantly increases the skill premium, from 1.60 to 2.51, leading enrollment to rise from 0.089 to 0.346. The mean earnings ratio rises to a lesser extent, from 1.89 to 2.44, as composition changes dampen the effect. The increase in the skill premium increases the cost of college relative to per capita output, c/y, by 30 percent. College costs themselves increase by almost $\$4,500$, as white-collar skill becomes more expensive.

Table 5. Effects of parameter changes

	y	c	c/y	MER	e	w^W/w^B
	(1)	(2)	(3)	(4)	(5)	(6)
1961 values: full model	29.0	9.80	0.338	2.17	0.201	1.00
2010 values: full model	64.9	26.93	0.415	2.44	0.346	2.51
Skill weight (ω_t)	70.8	22.47	0.317	1.89	0.089	1.60
TFP (A_t)	37.9	16.89	0.445	2.48	0.349	2.53
Co-pay rate (d_t)	64.9	26.94	0.415	2.44	0.348	2.51
Life-cycle earnings ($\varepsilon^B_{a,t}, \varepsilon^W_{a,t}$)	65.6	24.29	0.370	2.38	0.361	2.03
Population ($N_{a,t}$)	64.4	26.93	0.418	2.43	0.344	2.51
Interest rate ($r_{a,t}$)	60.8	25.53	0.420	2.45	0.343	2.52

Note:

y: GDP per capita, thousands of \$2005

c: College expenditures per student, thousands of \$2005

MER: Mean earnings ratio, defined in equation (19).

e: fraction of 18-year-olds enrolled in (and completing) college

w^W/w^B: Skill premium = relative wages, normalized by 1961 value.

In contrast, as the TFP parameter A increases from its 1961 value, the cost of college falls relative to output, from 0.445 to 0.415, because output grows more quickly than college expenditures. However, TFP growth generates the largest absolute increase in college costs, \$10,000 out of \$17,000, because it generates the largest increase in wages and GDP. TFP growth thus affects expected earnings, as workers expect skill prices to rise over their lives. (The steady state is really a balanced growth path.) TFP growth effectively steepens life-cycle earnings profiles. Because blue- and white-collar earnings profiles have different age slopes, and blue- and white-collar wages grow at different rates, it follows that TFP growth changes the relative lifetime earnings of blue- and white-collar workers associated with any sequence of wages. In the early years of our model, enrollment increases. Over time the stock of white-collar skill, W, grows relative to the stock of blue-collar skill, B, reducing white-collar wages. The earnings profiles shift as well. By 2010 the cumulative effect is a reduction in current enrollment, the mean earnings ratio and the skill premium.

Tuition as a fraction of cost, d, rises and falls several times between 1961 and 2010, making interpretation at any single date difficult. Allowing d to depart from its 1961 value reduces enrollment in 2010, but by only a small amount; we will return to this point below.

As Figure 8 shows, the changes in the efficiency profiles $\{\varepsilon_a^W, \varepsilon_a^B\}$ between 1961 and 2010 increased the stock of blue-collar skill (relative to white-collar skill) at any level of attainment, raising the college premium. College costs increase by over $\$2,500$ and also increase relative to output. On the other hand, the newer efficiency profiles imply lower earnings at younger ages and higher earnings at older ages. This decreases the discounted value of a college education, and enrollment decreases from 0.361 to 0.346. The baby boom, captured by changes in $\{N_a\}$, affects the composition of the blue-collar and white-collar workforces as well as their sizes. The demographic transition leads to a small decrease in the skill premium at 2010, but small increases in the mean earnings ratio and enrollment. Finally, interest rates have fluctuated over the past 50 years, with a slight decrease overall. Cheaper access to capital lowers the relative cost of college, from 0.420 to 0.415, leading to higher enrollment.

5.2 Higher costs and enrollment

It is commonly believed that the ongoing increase in college costs – and by extension tuition – has depressed college enrollment. To test this hypothesis, we fix college expenditures at their 1961 levels and recompute equilibrium enrollment, keeping tuition

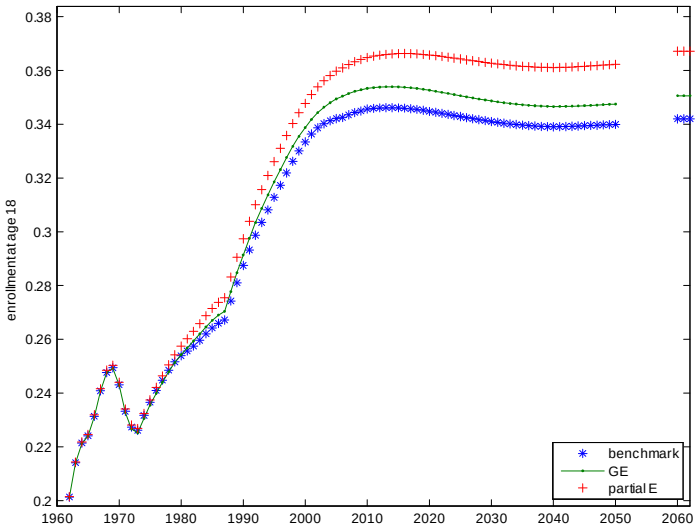

Figure 12: Enrollment when college expenditures are fixed at 1961 values: Baseline

discounts at their baseline values. Figure 12 shows the results. The blue line in this figure shows the benchmark enrollment trajectory, while the green line shows the enrollment trajectory that emerges when college costs are held fixed. By 2010, holding costs fixed increases enrollment by about 1 percentage point. The effect is small in large part because of general equilibrium effects. Any increase in enrollment also increases the relative supply of white collar skill, lowering the returns to college. As Heckman, et al. (1998b) note, such changes in wages, which affect earnings over an individual's entire working life, almost completely offset the changes in net tuition. The red line shows the enrollment that emerges when costs are fixed and wages are not allowed to adjust. Figure 12 shows that the partial equilibrium effect in 2010 is an enrollment increase of 2 percentage points. This is similar to the effect of temporarily reducing tuition to its 1961 levels for the years 2010-13; events that affect a subset of the population have small general equilibrium effects. In interpreting the modest drop in enrollment, it bears remembering that net tuition is less than 25% of total higher education expenditures, so that even the drastic cost reduction considered here reduces real net tuition by only about $\$4,000$.[19]

We repeat the experiment for the model with no borrowing constraints.[20] Figure 13 shows that in the absence of borrowing constraints, reducing costs to 1961 levels increases enrollment by a smaller amount. Figures 12 and 13 thus suggest that borrowing constraints do limit access to higher education, even if only modestly.

[19] See the first two rows of column (2) in Table 5.

[20] This version of the model also lacks induction costs, but the effects of that change are much smaller.

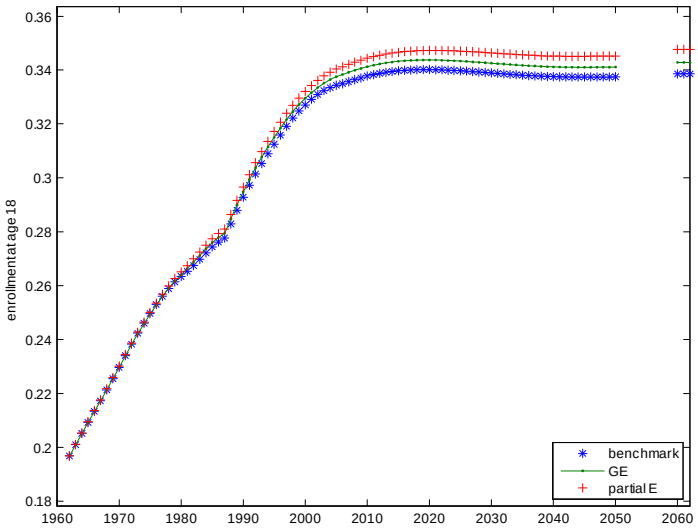

Figure 13: Enrollment when college expenditures are fixed at 1961 values: No borrowing constraints

5.3 Tuition discounts

In the final set of experiments, we estimate the dynamic effects of increasing the tuition co-pay rate (d_t). In the first set of experiments, we increase the co-pay rate by 10 percentage points over the years 1961-64. Figure 14 shows impulse responses for this change in a partial equilibrium setting, where wages are held fixed. The total tuition increase is biggest for the 1961 cohort ($t = 1$), which sees its net tuition go up by roughly 50% in all four years of college. Enrollment drops by more than 0.6 percentage points; the resulting elasticity is 0.073. The next three cohorts suffer price increases for decreasing portions of their college tenure, and their enrollment drops by decreasing amounts. After 4 years, tuition and enrollment return to normal. The effects of enrollment on the stocks of blue- and white-collar skill, are longer lasting, as the less-educated cohorts go through their working lives.

The partial equilibrium elasticity of 0.073 is comparable to a number of empirical estimates based on micro-level interventions (see Dynarski, 2003, and the papers referenced therein). Dynarski (2003, p. 285) finds that the elasticity of college attendance to total "schooling costs" is 1.5. Because most of this cost total consists of forgone earnings, Dynarski's estimates imply a tuition elasticity of 0.14.[21] While this is bigger than our

[21] Dynarski (p. 285) finds tuition and fees to be \$1,900, while foregone earnings are \$18,500, implying that her tuition elasticity is $1.5 \times (1900/(1900 + 18500))$. She also finds that each additional \$1000 of financial aid raises enrollment by 3.6 percentage points. After the financial aid program that she studies was eliminated, 35.2% of the (former) potential beneficiaries enrolled in college (p. 282). This yields an

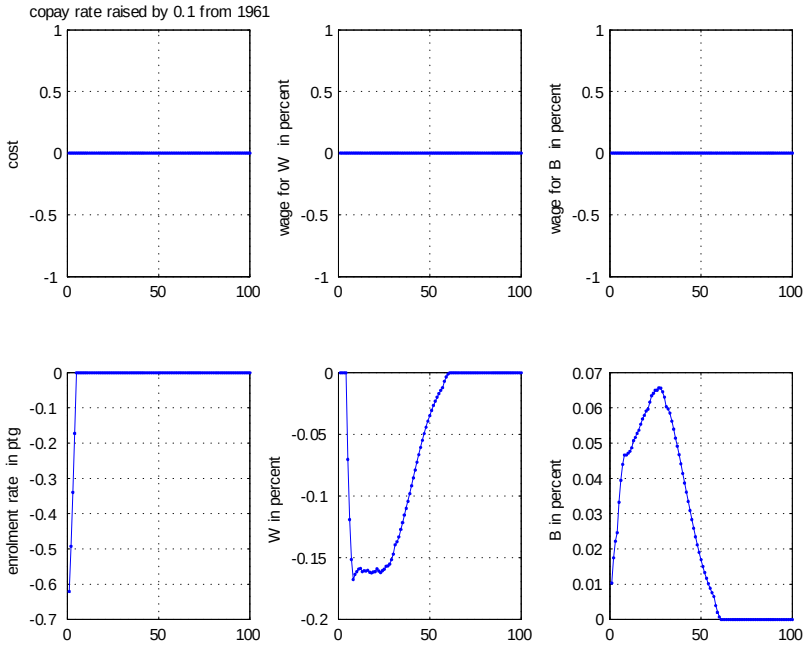

Figure 14: Effects of a temporary increase in d_t, partial equilibrium

estimate of 0.073, Dynarski is measuring the elasticity of college attendance, rather than earning a bachelor's degree – in any event, her estimated elasticity, like ours, is small.[22] Our estimated elasticity is also similar to that of Athreya and Eberly (2013). In a model where potential students hold limited financial wealth and face both drop-out and earnings risk, they find a partial equilibrium elasticity of 0.089.[23]

Figure 15 shows impulse responses in general equilibrium. In general equilibrium, the enrollment drop triggered by the tuition increase leads to an increase in white collar wages. This encourages enrollment by the cohorts entering immediately after 1964. Given the short duration of the tuition increase, however, the wage increase is small, and the enrollment elasticity for the 1961 cohort, 0.068, is very similar to the partial equilibrium elasticity.

Next, we consider a permanent 10 percentage point increase in tuition. Figure 16 shows the impulse response functions under partial equilibrium. The enrollment decrease

alternative tuition elasticity of $\ln([0.352 + 0.036]/0.352)/\ln([1900 - 1000]/1900) = -0.13$.

[22] While not comparing elasticities, Dynarski argues that her findings are comparable with a number of other studies, all of which find that "a $1,000 drop in schooling costs increases attendence by 3 to 4 percentage points" (p. 286). It bears noting that many of these studies are set in earlier years, where $1,000 represents a larger fraction of tuition than it does today.

[23] Athreya and Eberly find that decreasing d from 0.575 to 0.425 increases enrollment from 73% to 75%, implying an elasticity of $-\ln(0.75/0.73)/\ln(0.425/0.545)$. (Rounding error will at most make this a change from 72.5% to 75.5%, implying an elasticity of 13%.).

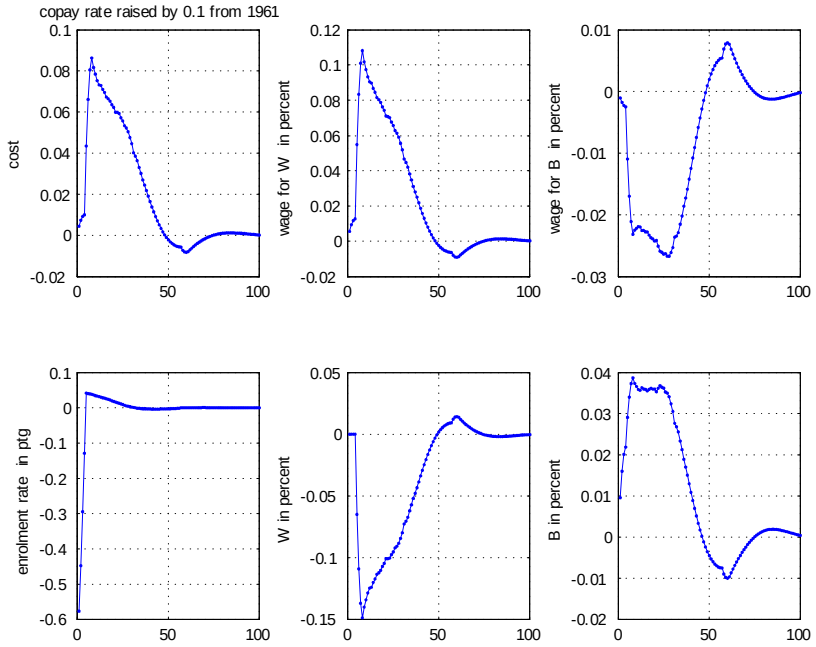

Figure 15: Effects of a temporary increase in d_t, general equilibrium

in 1961 – and the associated elasticity – is the same as for the temporary tuition decrease, but now it is followed by even larger decreases. These changes lead to significant changes in the stock of blue- and white-collar skill. Figure 17 shows the impulse responses under general equilibrium. The increase in white-collar wages generated by the decline in enrollment significantly reduces the size of the decline. Enrollment by the 1961 cohort drops by less than 0.35 percentage points, implying an elasticity of about 0.041. As the tuition increase continues, and the stock of white-collar skill continues to decline, the general equilibrium effect offsets the tuition increase to an even greater extent.

Figure 17 shows that the increase in white-collar wages leads to a similar increase in college costs; the effect on costs is a little smaller because of capital-labor substituion. This cost increase increases net tuition, reinforcing the effect of smaller tuition discounts. To assess the importance of this mechanism, we found the impulse response functions that arise when wages are allowed to adjust, but college costs are held fixed at their benchmark levels. Comparing these functions to those in Figure 17 shows that endogenous tuition provides almost no amplification of the original tuition shock. While a 10 percentage point increase in d represents a roughly 50% increase in net tuition, college expenditures increase by only about 0.7%; see the upper left panel of Figure 17. Even if the entire increase is passed on the student, net tuition rises by less than 3%. Recall that we have set the labor substitutability parameter ς to 0.7, following Heckman et al. (1998a). As

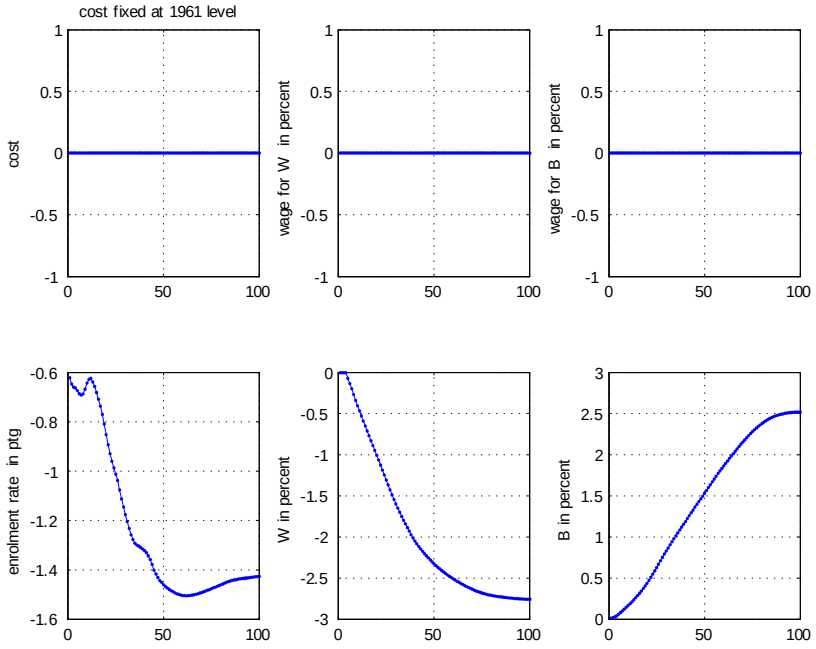

Figure 16: Effects of a permanent increase in d_t, partial equilibrium

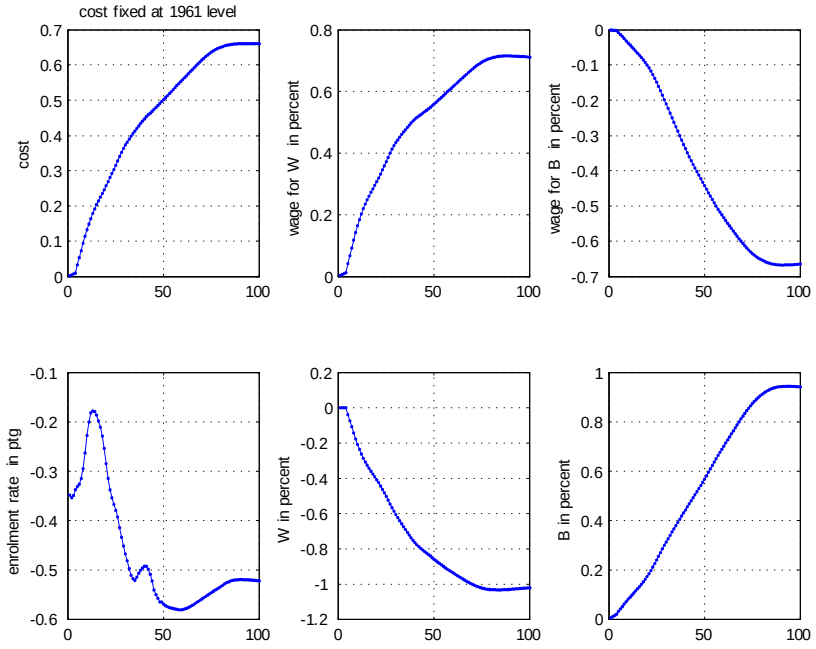

Figure 17: Effects of a permanent increase in d_t, general equilibrium

a result, changes in white-collar skill lead to only modest changes in white-collar wages and college costs. Unless this substitution elasticity is decreased significantly, endogenous tuition can not generate quantitatively meaningful amplification.

6 Conclusion

In this paper, we show that a general equilibrium model with skill- and sector-biased technological change can explain the increase in college costs observed over the past 50 years, along with the increase in college attainment and the increase in the relative earnings of college graduates. Our model has two key features. The first is the assumption that colleges have a limited ability to replace labor inputs with capital. In spite of the attention given to new education technologies, such as Massive Online Open Courses (MOOCs), this assumption appears consistent with historical experience. The second assumption is skill-biased technological change in the non-education sector. Together, these assumptions imply that skill-biased technological change makes a college education more expensive as well as more valuable.

We then use the model to perform a number of quantitative experiments. We find that the observed increase in college enrollment and the college premium are principally due to skill-biased technological change. Skill-biased change causes higher education costs to grow more quickly than GDP, while TFP growth in the non-education sector causes them to grow more slowly. Nonetheless, TFP growth generates the majority of the cost increase, because it generates the majority of the growth in wages. Our model suggests that if college costs had ceased to grow after 1961, enrollment in 2010 would be 1 to 2 percentage points (3 to 6 percent) higher. This effect is modest, but the tuition elasticities implied by our model are consistent with a number of previous studies. In our model, any event that reduces enrollment should, by raising white-collar wages, raise tuition in a way that amplifies the initial effect. The quantitative magnitude of this mechanism, however, turns out to be relatively small.

There are a number of useful ways to extend the model, including the introduction of heterogeneity in higher education. It would also be interesting to use our model for cross-country comparisons, or to consider primary and secondary education We leave all these topics for future work.

7 Appendix A: Data Construction

7.1 Enrollment

To construct our estimates of expenditures and tuition per-student, we need a measure of student enrollment that is concurrent with our measures of costs and revenues. The *Digest of Education Statistics* reports full-time equivalent (FTE) fall enrollment from 1967 to present. (FTE for for-profit institutions is provided only from 1990 on.) Total fall enrollment, which weights full-time and part-time students the same, is provided from 1963 to present, and for selected years back to 1869-70. Our preferred measure is FTE enrollment. Between 1949 and 1966 FTE enrollment for all private and all public institutions is found by multiplying total enrollment in each group by the ratio of FTE to total enrollment for 1967 for each group. Prior to 1949 we calculate FTE enrollment only for all institutions, as total enrollment multiplied by the ratio of FTE to total enrollment in 1949. Prior to 1967, public FTE enrollment is divided between 2-year and 4-year institutions on the basis of total enrollment. Between 1967 and 1989, private FTE enrollment is divided between for-profit and not-for profit institutions on the basis of total fall enrollment. Prior to 1967, we assume that enrollment is split between for-profit and not-for-profit institutions at its 1967 proportions.

This measure of enrollment should be distinguished from our measure of college attainment, described below, which we construct from the Current Population Survey. The two measures differ conceptually because the attainment measure treats people who earned Bachelors degrees in their late 20s or early 30s as having attended college immediately after high school; this is consistent with our model, where college must be attended without delay. The enrollment measure makes no such adjustments.

7.2 Educational Expenditures and Gross Tuition

Educational and general (E&G) expenditures are a category of the operating cost data presented in the *Digest of Education Statistics*. Among the costs excluded from this category are auxiliary operations such as dormitories and hospitals. As discussed further below, these data also exclude most interest expenditures, which are not considered operating expenses. All the cost (and revenue) data are measured over academic years, July 1 - June 30, which we index by the initial calendar year.

As described on Department of Education's website (U.S. Department of Education, 2012a), the accounting methods used to construct these data have changed over time.

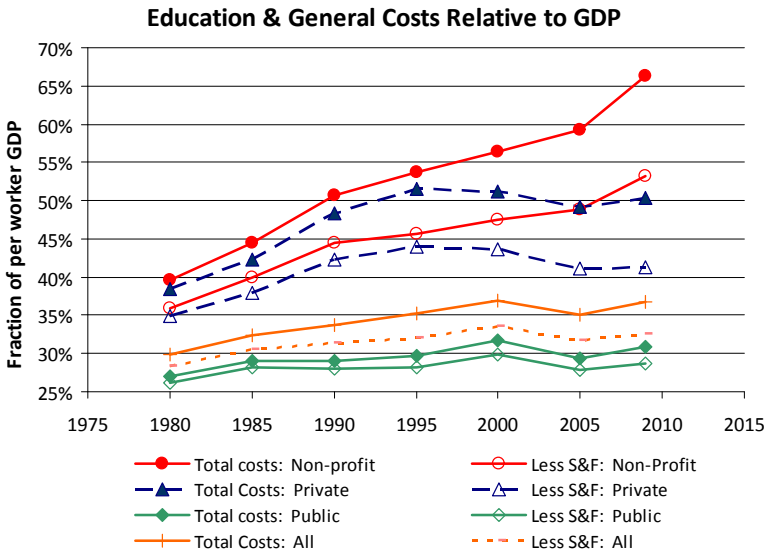

Figure 18

Through 1995-96, all institutions used the "old form". Private institutions adopted a new accounting methodology in 1996-97, while public institutions continued to use the old form until 2000-01, before adopting a third accounting methodology in 2003-04.

One quirk of the old form data is that E&G expenditures included institutional "scholarships and fellowships": rather than being deducted from tuition revenue, this aid was treated as an expense. While the current methodology for public institutions still treats institutional aid as an expense, the methodology for private institutions now deducts "institutional tuition allowances" from tuition revenues, and reduces scholarship and fellowship expenses accordingly. The residual scholarship and fellowship expenditures are now quite small.

Our approach is to include all scholarships and fellowships in our measure of "sticker price" tuition, deduct them (as institutional grant aid) from our measure of net tuition, and exclude them entirely from E&G costs – institutional aid is a discount in price, rather than an purchase of resources. For private sector institutions in the post-old-form era, we download institutional tuition allowances from the Department of Education's IPEDS data center (U.S. Department of Education, 2012b), and add them to tuition revenue.[24]

Figure 18 shows the effects of excluding scholarships and fellowships from our definition of E&G expenditures. We impute scholarships and fellowships for private institutions in the post-old-form era by multiplying institutional grant aid (also downloaded from the

[24] We are grateful to Colleen Lenihan for assisting us with these data.

IPEDS data center) by the ratio of scholarship and fellowship costs to institutional grant aid (from the College Board, 2011) in 1996. Figure 18 reveals that scholarships and fellowships appear to have become increasingly important over time. In 2009, they equalled 4 percent of per worker GDP.

E&G expenses are not explicitly calculated in the post-old form data. For the public sector, we calculate E&G expenses by first finding the sum of instructional, research, public service, academic support, student services, institutional support, operation and maintenance, and scholarships and fellowships (when included). These categories were the principal categories of E&G expenditures in the old form data. (The old form data also include mandatory transfers.) We then add a portion of depreciation and interest expenses, which we allocate proportionally between (non-scholarship) E&G expenses and auxiliary uses. This creates a discrepancy with the old form data, which did not include depreciation. The old form data did include expenditures on new equipment and furniture—the new data do not—but these expenses were not listed separately. A second discrepancy arises because the old form data did not include most interest expenses either, although it did include some repayment of principal. (See U.S. Department of Education, 2012a.) In addition, the old form data do not list interest expenses separately. Fortunately, interest expenditures appear to be less than 2 percent of total operating expenses. Moreover, under both accounting systems, interest expenditures are measured using an accounting cost, rather than an opportunity cost, approach. The latter approach, where interest expenses are based on the market value of all assets, rather than the book value of outstanding debt, should yield higher costs.

We calculate E&G expenses for private institutions in a similar way. In the post-old form private sector data, depreciation expenses are included in the functional subtotals (e.g., instruction or dormitories). The functional subtotals also include operation and maintenance (O&M) expenditures, which are listed separately and treated as E&G expenses in both the old form data and the post-old form public institution data. We impute the O&M expenditures associated with auxiliary functions on the basis of private institution data for 1995-96 (which gives us O&M as a fraction of non-scholarship operating expenses), and add it to our estimate of private sector E&G costs. We are unable to identify interest expenses.

The Department of Education provides separate cost data for 4-year public institutions dating back to 1976-77. Although separate costs for 4-year private institutions go back only to 1996-97, a much smaller fraction of private institutions are 2-year. Combining the 4-year public institutions with all the private institutions produces an approximate

set of 4-year institutions. Prior to 1976, we allocate costs between the 2- and 4-year of institutions as follows. As a matter of notation, costs for all institutions in year t can be written as $C_t = C_{2t} + C_{4t}$, where "2" and "4" denote subtotals for 2-year and 4-year institutions, respectively. Likewise, enrollment can be written as $E_t = E_{2t} + E_{4t}$. Finally, let $c_{2t} = C_{2t}/E_{2t}$ and $c_{4t} = C_{4t}/E_{4t}$ denote per-student expenditures, and let $x_t = c_{2t}/c_{4t}$ denote their ratio. As a matter of algebra, it can be shown that $c_{4t} = C_t/[E_{4t} + E_{2t}x_t]$, so that $C_{4t} = C_t/[1 + x_t \cdot E_{2t}/E_{4t}]$. Combining the 1976-77 value of x_t with enrollment data allows us to estimate cost totals for 4-year institutions in earlier years. In short, by assuming that the expenditures per 2-year student are a constant fraction of expenditures per 4-year student, we can use enrollment data to impute disaggregated costs. Castro and Coen-Pirani (2012) adopt a similar approach. We also use this approach to allocate costs between private not-for-profit and for-profit institutions prior to 1990, and (by reversal) to impute private for-profit expenditures and revenues between 1996 and 1998. In the latter case, we use weighted averages of x_t immediately before and after the imputation years.

In spite of the many imputations, we find that the our measure of E&G costs, as a fraction of per capita GDP, evolves smoothly during the accounting transition between 1995 and 1997. While there is missing data in 2001 and 2002, the cost ratios in 2000 and 2003 are similar.

In years where no expenditure data is available at any level, we assume that the nominal cost to GDP ratio evolves linearly, and multiply the imputed ratio by GDP. 4 of the 49 values used as estimation data are imputed this way.

Our measure of gross tuition is found by dividing tuition revenues (including imputed tuition allowances) by FTE. This measure is consistent with our measure of costs, and it can be extrapolated back to 1929. On the other hand, this measure combines both undergraduate and graduate tuition. Since 1964, the Department of Education has estimated average undergraduate tuition, and since 1977, it has estimated average undergraduate tuition at 4-year institutions. Figure 19 compares the two measures of tuition. Over the period 1964-2009, the two sets of measures imply similar changes in tuition.

For missing values, we impute tuition per 4-year student under the assumption of geometric growth. While we use a shorter sample, the disaggregated tuition data also has a number of missing values. When possible we impute them using the same methodology as for costs.

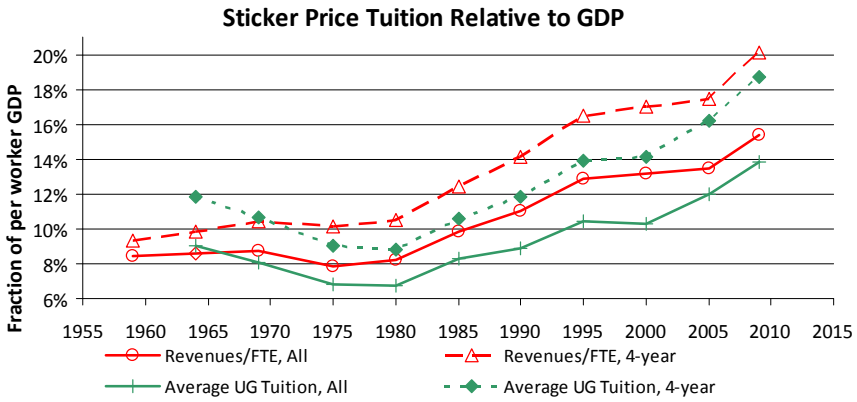

Figure 19

7.3 Financial Aid

Our measure of financial aid comes from two sources. We measure institutional financial aid per FTE using the (adjusted) scholarships and fellowship expenditures described above. We measure non-institutional aid using data provided by the College Board (2011). These data list nominal grant aid by source (Table 2), which we sum. We convert these totals into per-FTE quantities using the College Board's enrollment measure (Table 4), which differs only slightly from those calculated by the Department of Education. Because the College Board does not provide estimates of FTE in 1970-71 or 1963-64, for those years we use the Department of Education quantities. For years where costs are missing, non-instituional grant aid per FTE is imputed under the assumption of constant exponential growth. The College Board data do not allow us to calculate aid quantities disaggregated by type of institution. Our measure of net tuition is found by deducting institutional and non-instituional financial aid from gross tuition.

Our measure of the tuition co-pay rate, d, is the ratio of tuition net of grant aid to E&G costs, all of which are calculated on a per-FTE basis. Our preferred measure of d is based on cost, tuition, and institutional aid data for our approximate set of 4-year institutions; our measure of non-institutional grant aid, however, is the same for all types of institutions. We assume that d remains at its 2009 value of 23.2% through 2300. Missing values for earlier years are linearly interpolated.

7.4 Demographics

We begin with Census data (inclusive of Armed Forces overseas) for the period 1961-2000. For the period 2000-2050, we use the Census Bureau's 2008 projections. These

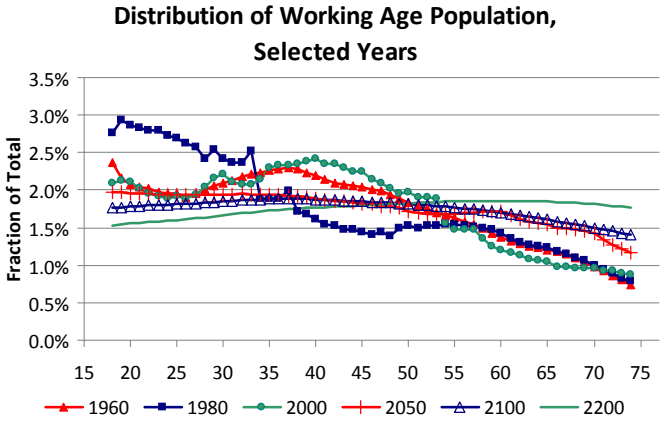

Figure 20

detailed projections allow us to calculate "survival" rates (net of immigration) at every age and the growth rate of the birth cohorts. We project all of these rates for the period 2051-2100 by linearly extrapolating their values over the period 2040-2050, and use the projected rates to update the population. From 2100 forward we assume that the size of the birth cohort is constant (in this respect, we follow Krueger and Ludwig, 2007), and that the net survival rates are constant as well. By 2200 the population has converged to a stationary distribution. Our calculations use the subset of the age distribution covering ages 18-74. To make the population data correspond with the academic year used by the Department of Education, we take averages of consecutive years; for example, the population for the 1965-66 academic year is found by averaging the population values for calendar years 1965 and 1966. Figure 20 shows how the age distribution of workers evolves over time, as the population ages.

7.5 GDP per capita

Our measure of GDP is taken from the National Accounts data assembled by the U.S. Bureau of Economic Analysis (BEA). Dividing GDP by the number of people aged 18-74 yields our measure of GDP per capita. To make these data correspond with the academic year used by the Department of Education, we take averages of consecutive years; for example, GDP per capita for the 1965-66 academic year is found by averaging over the calendar years 1965 and 1966. The BEA provides both nominal and real measures of GDP, with the real data measured in year-2005 dollars. Dividing nominal quantities by real produces an "academic year" GDP deflator.

7.6 Capital

The Digest of Education Statistics reports total physical plant in the higher education sector for selected academic years between 1899-1900 through 1994-95. These measures are based on book, or historical, costs rather than market value. We convert book values into current cost (market) values by multiplying them by the ratio of current cost to historical cost capital found in the BEA's fixed asset data (including government assets). We make separate adjustments for equipment and structures when possible. In years where the Digest reports only total book costs, we rescale using total cost ratios from the BEA, adjusted by the aggregation error observed in years where all data are available.

A major limitation of the capital data reported in the Digest is that they do not distinguish capital devoted to "education and general" uses from capital devoted to auxiliary "current fund" uses such as dormitories. We allocate capital between uses by assuming that capital is proportional to operating costs (exclusive of scholarships and fellowships).

To compute aggregate capital-output ratios, we combine the estimate of total fixed assets provided by the BEA with the BEA's measure of GDP.

The capital measures provided in the Digest, like those produced by the BEA, are end-of-period quantities. Because we measure capital in our model as a beginning-of-period quantity, we measure capital as the previous period's end-of-period value. In earlier years where the capital stock is reported only occasionally, this produces mis-alignments with the operating cost and enrollment data. We impute capital for these missing periods under the assumption of geometric growth.

7.7 College Attainment

Our estimates of college attainment presented in Section 2.4, the fraction of the population with at least a Bachelors degree, are drawn from the IPUMS implementation of the Current Population Survey. We work with data from the annual March survey. Because most people have completed their undergraduate studies by the time they reach 30, we measure each cohort's attainment as the average across ages 28-32. Assuming the attainment rates have a quadratic trend and an autoregressive residual, we use the data for 1936-1979 to project attainment for people born between 1980 and 1990.

The educational attainment of people 19 and older in 1961 – cohorts born before 1943 – is also taken from the CPS. When possible, we measure educational attainment as the average over ages 28-32; for older cohorts we take averages across the 5 (if available) youngest ages observed in our data.

7.8 Earnings

Our estimates of the wages and earnings shown in Section 2.4 are also calculated from IPUMS CPS data. Our construction of these variables, and especially the top-coding adjustment for income, most closely follows the approach in Heathcote, Perri and Violante (2011). In calculating hours (for constructing wages), we convert interval responses into numerical values with the approach described in Abraham, Spletzer and Stewart (1998). Sorting people by their current education level, we calculate average earnings for people whose highest education level is high school and those with at least a Bachelors degree. This allows us to calculate mean earnings ratios. We linearly impute the ratio for the missing 1962 data.

7.9 Life Cycle Efficiency Profiles

We estimate the efficiency profiles $\{\varepsilon^W_{a,t}, \varepsilon^B_{a,t}\}$ from CPS data on earnings. Suppose that agent i is age a_{it} at time t. It follows that the agent's "birth year", b_i, equals $t - a_{it}$. Let $S_i \in \{B, W\}$ index the agent's education level. As described in the main text, we assume that log earnings, $\ln(\widetilde{y}^S_{a,t})$, follow.

$$\begin{aligned} \ln(\widetilde{y}^S_{a,t}) &= \xi^S_0 + \xi^S_1 a + \xi^S_2 a^2 + \xi^S_3 a^3 + \xi^S_4 t \cdot a + \xi^S_5 t \cdot a^2 + \xi^S_6 t_2 \cdot a + \xi^S_7 t_2 \cdot a^2 \\ &\quad + \zeta^S_1 e_b + \zeta^S_2 e_b^2 + \zeta^S_3 e_b^3 + \zeta^S_4 e_b^4 + \sum\nolimits_u \lambda^S_u \cdot 1\{t = u\} + \eta^S_{a,t}, \end{aligned} \tag{21}$$

where: $\eta^S_{a,t}$ is classical measurement error; e_b denotes the fraction of cohort b with a college degree; $1\{t = u\}$ is a time dummy; and $t_2 = \max\{t - 1990, 0\}$ is a spline term. Because we are interested in overall earnings, including those of non-workers, we estimate equation (21) on a synthetic panel (Deaton, 1985): each value of $\ln(\widetilde{y}^S_{a,t})$ in our estimation dataset is the log of average earnings for all individuals—working or not—of age a in calendar year t.

Table 6. Coefficient Estimates for Earnings Profiles

Variable	High School	Bachelors Degree+
Constant	9.772 (0.038)	10.586 (0.043)
Age	−3.605 (0.320)	−1.247 (0.360)
Age^2/100	9.228 (0.593)	4.619 (0.744)
Age^3/10000	−0.469 (0.012)	−0.540 (0.017)
Year×age/1000	1.831 (0.162)	0.643 (0.182)
Year×age^2/100000	−4.594 (0.300)	−2.268 (0.377)
max{year−1990,0}×age/1000	−2.782 (0.338)	−2.037 (0.430)
max{year−1990,0}×age^2/100000	8.261 (0.583)	6.353 (0.782)
Fraction with Bachelor's degree	4.796 (4.419)	19.81 (4.913)
(Bachelor's fraction)^2	−68.82 (37.46)	−178.2 (42.38)
(Bachelor's fraction)^3	290.7 (134.8)	636.8 (153.5)
(Bachelor's fraction)^4	−386.0 (173..3)	−803.3 (197.8)
R^2 (with dummies)	0.939	0.917
Average time dummy	−0.0003	0.0004
R^2 (without dummies and fractions)	0.906	0.878

Notes:

1. Age for High School regression = age - 20.
2. Age for Bachelors Degree regression = age - 25.
3. Coefficients for individual time dummies not shown.

We assume that $\ln\left(\varepsilon^S_{a,t}\right)$ equals the right-hand-side terms on the first line of equation (21); the terms on the bottom line reflect selection effects and factor prices that are produced within the model. We normalize the attainment terms on the bottom line to have a mean of zero, and add the mean time effect to ξ^S_0. This is only a normalization: as we explain in the main text, it is the slopes of the profiles, not their absolute levels, that are the key inputs to our model. The coefficients $\{\xi^S_4, \xi^S_5, \xi^S_6, \xi^S_7\}$ capture date-specific trends in life cycle labor supply and human capital accumulation. We treat these trends as exogenous and include them in the profiles that we feed in our model. Table 6 shows

our estimation results. When solving the model we assume that the efficiency profiles evolve in this fashion until the 1970 cohort is reached, after which they remain constant.

The time dummies, not shown here, imply that the market price of white-collar labor has been rising over the sample period, while the price of blue-collar labor has been falling. (See Figure 11.) Evaluating the polynomials in the attainment fraction (using the data in Figure 6) shows that the average quality of white-collar workers has fallen, consistent with standard selection arguments. The quality adjustment for high school graduates, on the other hand, shows no long term trend.

8 Appendix B: Standard Errors

We are matching summary statistics—GDP, attainment rates, the wage premium, higher education expenditures, and higher education capital—for the years 1961-2009. Let f_{mt}, $m \in \{1, 2, ..., M\}$, $t \in \{1, 2, ..., T\}$, denote an observation of type m in year t. In most cases, we only have f_{mt} itself, not the disaggregated underlying data used to construct it. However, each f_{mt} can be considered a time series realization.

The model-predicted value of f_{mt} is $f^*_{mt}(\theta)$, where $\theta \in \mathbb{R}^J$ is the parameter vector we wish to estimate. Let θ_0 denote the population value of θ. We assume that the normalized deviation, $u_{mt} = f^*_{mt}(\theta_0)/f_{mt} - 1$, is zero-mean conditional on $f^*_{mt}(\theta_0)$, stationary and ergodic. Variation in u_{mt} can be due to measurement error in f_{mt}, transitory shocks to $\{A_t, \omega_t\}$, or other shocks not included in our model.

Our estimate of θ_0, $\widehat{\theta}$, solves

$$\min_{\theta} \sum_{t=1}^{T} \sum_{m=1}^{M} \left(\frac{f^*_{mt}(\theta)}{f_{mt}} - 1 \right)^2.$$

The first-order conditions are

$$\sum_t \sum_m u_{mt}(\widehat{\theta}) \frac{1}{x_{mt}} \frac{\partial f^*_{mt}(\widehat{\theta})}{\partial \theta_j} = 0, \quad j \in \{1, 2, ..., J\}.$$

These equations provide a set of moment conditions compatible with the Generalized Method of Moments approach. Defining the $M \times 1$ vectors $f^*_t(\theta) = [f^*_{1t}(\theta), ..., f^*_{Mt}(\theta)]'$, $u_t = [u_{1t}, ..., u_{Mt}]'$, $f_t^{-1} = [1/f_{1t}, ..., 1/f_{Mt}]'$, we can rewrite the J first-order conditions as

$$\sum_t \left[\frac{\partial f^*_t(\widehat{\theta})}{\partial \theta'} \right]' [u_t(\widehat{\theta}) \odot f_t^{-1}] \equiv \sum_t g_t(\widehat{\theta}) = \underset{J \times 1}{0},$$

with "$\odot$" denoting element-by-element multiplication.

It follows from standard arguments that under regularity conditions

$$\sqrt{T}(\widehat{\theta} - \theta_0) \xrightarrow{d} \mathcal{N}(0, D^{-1}S(D^{-1})'),$$

where

$$D = \frac{\partial}{\partial \theta'} E\left(g_t(\theta_0)\right)$$

is the $J \times J$ gradient matrix of the first-order conditions and

$$S = \sum_{j=-\infty}^{\infty} E\left(g_t(\theta_0) g_{t-j}(\theta_0)'\right)$$

is the spectral density matrix of the moment vector.

We approximate D with

$$\widehat{D} = \frac{1}{T} \sum_t \frac{\partial g_t(\widehat{\theta})}{\partial \theta'},$$

and S with a Newey-West estimator.

Ideally the standard errors would account for uncertainty in the first-step parameters. Given the broadness of our first-step calibration, however, such an adjustment is not feasible. An additional complication is that we do not observe educational capital for the entire sample period. To account for this possibility, we define $u_{mt}(\widehat{\theta}) \frac{1}{f_{mt}} \frac{\partial f^*_{mt}(\widehat{\theta})}{\partial \theta_j}$ to equal zero when the capital data are missing. Under the assumption that the pattern of observed/missing data is "scaled up" as the sample grows, our asymptotic results still hold under this modification.

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