Wage Structure and Public Sector Employment: Sweden versus the United States 1970-2002

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Abstract

Swedish census data and tax records reveal an astonishing wage compression; the Swedish skill premium fell by more than 30 percent between 1970 and 1990 while the U.S. skill premium, after an initial decline in the 1970s, rose by 8–10 percent. Since then both skill premia have increased by around 10 percentage points in 2002. Theories that equalize wages with marginal products can rationalize these disparate outcomes when we replace commonly used measures of total labor supplies by private sector employment. Our analysis suggests that the dramatic decline of the skill premium in Sweden is the result of an expanding public sector that today comprises roughly one third of the labor force, and that expansion has largely taken the form of drawing low-skilled workers into local government jobs that service the welfare state.

Key words: Skill premium, employment, private sector, public sector, Sweden, United States.

JEL-classification: E24, J31.

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1 Introduction

The last three decades have seen a sharp increase in the skill premium for college educated in the United States while workers in Europe have experienced a compression of wage structures with lower skill premia, which is perhaps starkest seen in Sweden.[1] According to Table 1, the U.S. skill premium rose by 8–10 percent between 1970 and 1990 and an additional 11 percentage points in 2002. During the same period 1970-1990, the Swedish skill premium fell by more than 30 percent and recovered by no more than 10 percentage points in 2002. These numbers are computed by dividing all employees into 'skilled' and 'unskilled' workers where skilled are those with a traditional college education (at least three years of university studies in Sweden and a total of at least 16 years of schooling in the U.S.). The Swedish data is taken from census data and tax records which include Sweden's entire population, while the U.S. data is drawn from the *Current Population Survey* (CPS).

Table 1: Skill premia and labor supplies in the U.S. and Sweden
(See Appendix A for details.)

	U.S.		Sweden	
	1990	2002	1990	2002
Private sector				
Skill premium relative to 1970	1.099	1.210	0.687	0.790
Ratio of skilled to unskilled labor, log change relative to 1970	0.82	1.08	1.24	1.68
Aggregate economy				
Skill premium relative to 1970	1.077	1.190	0.668	0.718
Ratio of skilled to unskilled labor, log change relative to 1970	0.66	0.90	0.86	1.22

These vastly different outcomes are not a result of any significant difference in initial absolute skill premia; Edin and Topel (1997, p. 173) report that the typical college graduate in 1968 earned 50% percent more than an otherwise comparable high school graduate in both the U.S. and Sweden.[2] This has led many observers to propose nonmarket explanations to the evolution

[1] For some evidence, see Freeman and Katz (1995).

[2] In a contemporary study of rates of return to education in Sweden, Ståhl (1974) reports examples of earnings differentials that are of the described magnitude; e.g., an average annual income of 85 000 SEK for 45-55 year old engineers and of 50 000 SEK for 50 year old blue-collar workers.

of the Swedish skill premium. For example, Edin and Topel suggest that the answer should be sought in a theory of collusion between the trade union and the employers' association.[3] Moreover, as Sweden is a small open economy, it is difficult to argue that Sweden has not been exposed to similar technological changes and forces of globalization as the U.S.

Thus, these dramatically different outcomes pose a challenge to economic analysis based on times series of aggregate labor supplies, such as Katz and Murphy (1992), hereafter denoted KM, and Krusell, Ohanian, Ríos-Rull and Violante (2000), hereafter denoted KORV, who study the historical U.S. skill premium. They show that the evolution of the U.S. skill premium can successfully be accounted for by observed variations in the relative supplies of skilled and unskilled labor. In this paper we ask whether variations in relative supplies of skilled and unskilled labor can account for the evolution of the Swedish skill premium under the assumption that Sweden has faced the same skill-biased technological change as estimated for the U.S.?

We find that aggregate Swedish relative supplies of skilled and unskilled labor cannot explain the observed dramatic decline in the skill premium. However, when we differentiate between public and private sector employment, relative supplies of skilled and unskilled labor can explain the observed dramatic shifts.

We adopt the equilibrium theory of wage determination that underlies the quantitative studies of KM and KORV, i.e., the skill premium is determined by the marginal products of skilled and unskilled labor; but we question their use of economy-wide labor supplies. Their implicit assumption is that total employment contributes to a composite output that is demanded by households and efficiently produced – an assumption that seems questionable regarding public sector employment. The public sector does not only produce public goods and services that differ from what would be demanded and produced, if the workers were instead employed in the private sector; but even when public sector output is a close substitute to private sector output, the public sector would most likely neither attain the efficiency nor reflect the scale and composition of an alternative market-driven allocation in the private sector.[4] To emphasize these differences, we assume that the private sector produces a composite output that is determined by market demand and that is distinct from public sector output. Hence, the proper objects of our study are not economy-wide labor aggregates but rather labor inputs in the private sector. This distinction turns out to be immaterial for the study of the U.S. skill premium but of critical importance when explaining Swedish labor market outcomes.

[3]Edin and Topel (1997, pp. 192–197) favor a theory in which the egalitarian objectives of the trade union coalesce with the interests of the employers' association when the bargained "agreements did not just raise the compensation of low-wage workers; they also reduced the absolute wage of skilled workers . . . delivering 'cheap' skilled labor to large employers."

[4]An example of differences in public and private sector allocations is offered by Rosen (1997), who studied the public provision of child-care services in Sweden. Rosen concludes that there is an excessive supply of such services in Sweden which "reduces the value of social output and living standards in the overall economy. Total output is smaller than it would have been if household services had been paid for privately and transacted through the market."

In contrast to the U.S., Sweden has seen a rapidly expanding public sector in the 1970s which has had a major impact on the supply of labor available to the private sector. Public sector employment as a fraction of all employed were rather similar in 1970; 22 percent in the U.S. and 24 percent in Sweden.[5] By the mid 1980s the countries looked very different. In the U.S., the share had increased by one or two percentage points. In Sweden on the other hand, public sector employment had increased by more than 50 percent and amounted to 35 percent of total employment.

The expansion of the public sector does not have a direct effect on the skill premium, as any economic model with free labor mobility implies that changes in the skill premium should be the same across sectors[6] – an implication that is also largely born out in the data of Table 1. Instead, the influence of the public sector upon the skill premium is an indirect one where public employment decisions affect the ratio of skilled to unskilled workers available to the private sector. The numbers and skill composition of public employees are not constrained by profit considerations in the market place, but rather guided by public policies and largely financed through taxation. Crucial to our theory is that the expansion of the Swedish public sector disproportionately involved the hiring of unskilled workers. This is reflected in Table 1 where the logarithmic increase in the ratio of skilled to unskilled labor was much larger in the private sector as compared to the aggregate Swedish economy. In the U.S. on the other hand, the changes in the ratio of skilled to unskilled labor were of more similar magnitudes in the private sector and in total employment.

We consider two different approaches in analyzing the relationship between changes in relative labor supplies and the evolution of the skill premium. First, we follow KM and assume that relative demand effects can be captured using a simple linear time trend. Second, we follow KORV and assume that relative demand effects are due to capital-skill complimentarity in combination with equipment-specific technological progress, as measured by quality-adjusted equipment price indexes. Under our auxiliary assumption that public sector employment does not substitute for private sector employment, we can then explain the disparate developments of skill premia in the U.S. and Sweden with both approaches, especially with the latter one. The dramatic decline of the skill premium in Sweden is the result of an expanding public sector that today comprises roughly one third of the labor force, and that expansion has largely taken the form of drawing low-skilled workers into local government jobs that service the welfare state.

Our findings suggest that the causes to compressed skill premia in other European countries might be sought in the withdrawal of low-skilled workers

[5]We define public sector employment as those employed in health, education, postal services and government administration. While health and to some extent education are largely produced in the private sector in the U.S., they are almost exclusively produced in the public sector in Sweden. To facilitate comparison between the two countries we choose to use the Swedish division into private and public employment for both countries. Though for the U.S., the distinction between private and public employment is immaterial for the results.

[6]In fact, economic models with free labor mobility imply that wages are equalized across sectors for the same type of labor. Our study does not address discrepancies in that dimension.

from the private sector into public sector employment, long-term unemployment, or disability and early retirement programs.

The paper is organized as follows. Section 2 describes the data. Section 3 contains the analysis using the approach of KM. Section 4 conducts the corresponding analysis using KORV's model, with a sensitivity analysis and a look at additional implications in Section 5. Section 6 concludes with a discussion of our findings. The appendices provide some further details on the data and KORV's framework.

2 Data

In this section we briefly describe the data and how we construct our measures for skill premia and labor inputs. Our procedure is very similar to that employed by KM and KORV. The wage measure used throughout the paper is average hourly wages. Labor input is measured in efficiency units. For a more complete description, see Appendix A.

2.1 U.S. data

The source for our U.S. data are the 1971-2003 March CPS Annual Demographic Survey files provided by UNICON, from which we extract data for the years 1970-2002. From these data we construct two different samples: (i) a supply sample including all workers between 16 and 70 years of age and (ii) a wage sample which is restricted to full-time workers. In each sample, we sort workers into groups according to age, sex, race, education, and private/public sector employment.

We treat individuals within groups as perfect substitutes. Using the wage sample we calculate with-in group average hourly wages as the ratio between total income and total annual hours for each group. Using the supply data we calculate total annual hours for each group.

The groups are then sorted into two classes; skilled and unskilled labor, where by skilled we mean college graduates. We obtain class-specific measures by aggregating across groups. Total labor supply for each class is given by weighting hours in each group within the class by average group wages and then summing. The average wage for each of the two skill classes is then calculated simply as the ratio between total class income and total class labor input.

2.2 Swedish data

We apply the same procedure on the Swedish data. The sources for our Swedish data are the Census of Population Surveys for 1970 and 1990, and the LOUISE database for the period 1990-2002. The Census of Population Surveys cover *all* Swedish individuals, and the LOUISE database covers all individuals between 16 and 64 years of age for the period 1990-95 and all individuals above age 16 thereafter. This is a remarkable combination of data sources that to our knowledge has not been employed previously. All income data is based on tax records, which implies that there are no problems associated with self-reporting or with topcoding.[7] Hence, for the years 1970 and 1990-2002, we obtain as close as possible to perfect measures of the skill premium and of the relative supply of skilled and unskilled labor. When discussing our findings, we will therefore emphasize these observations.

For the years 1971 through 1989 there exist no comparable data source in Sweden. To trace out the development of the skill premium for these years we rely on results from a study by Edin and Holmlund (1995, table 9.2). Using survey data they estimate year-specific human-capital type wage equations for several years between 1968 and 1991.[8] We assume that our measure of the skill premium has the same time profile as Edin and Holmlund's estimated wage differential between workers with 16 and 12 years of education. It is comforting to note that the change in their estimated wage differential between 1968 and 1991 is very similar to the change in the skill premia we derive using the Census data from 1970 and 1990.

To trace out the development of labor inputs for the years 1971 through 1989, we rely on data from labor-market surveys (*Arbetskraftsundersökning*, AKU) that also gather information on workers' educational attainment and industry classification.[9] For these years we assume that our measure of labor input has the same time profile as the ratio of number of skilled to unskilled in the survey data. It is once again comforting to note that the change in the ratio of skilled to unskilled in the survey data over the period 1971 (the first year for which we have survey data) and 1990 is very similar to the change in our measure of labor input using the Census data from 1970 and 1990.

2.3 Trends in labor supply and skill premia

We lay out the basic facts in Figures 1 and 2 that describe the private sector. Figure 1 shows a substantial rise in the relative supply of college educated workers in both Sweden and the U.S. The relative supply of skilled workers in the U.S. rose by 193 percent between 1970 and 2002. The percentage change is even larger in Sweden where the supply of skilled workers rose by 435 percent between 1970 and 2002.

Figure 2 shows how the Swedish and U.S. skill premia evolved between 1970 and 2002. The U.S. development is well known. The skill premia deteriorated during the 1970s, but has grown rapidly since around 1980.[10] The development of the Swedish skill premia is dramatically different. Throughout the period the

[7] See Appendix A.4 for a discussion of topcoding in the U.S. data.

[8] Their data source is the Level of Living Survey for 1968, 1974 and 1981 and the Household Market and Nonmarket Activity Survey for 1984, 1986, 1988 and 1991.

[9] We thank Per-Anders Edin for providing us with the data. For further details on the data and our interpolation, see Appendix A.5.

[10] Our time series for the skill premium in the U.S. is very similar to that documented by Autor et al. (2005). Recently, some authors have argued that the rise in skill premia during the 1980s was an episodic event that did not continue in the 1990s. However, as shown in Appendix A.4, this argument relies to a large extent on how the topcoding is done.

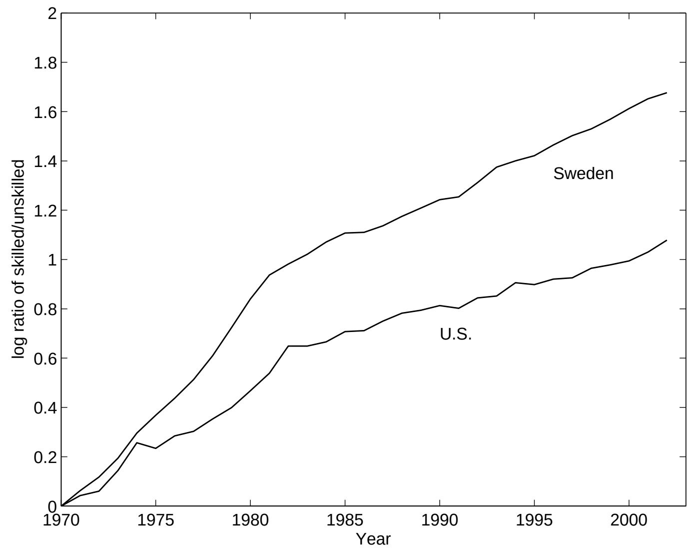

Figure 1: Ratio of skilled to unskilled labor in Sweden and the U.S. in the private sector (log change relative to 1970).

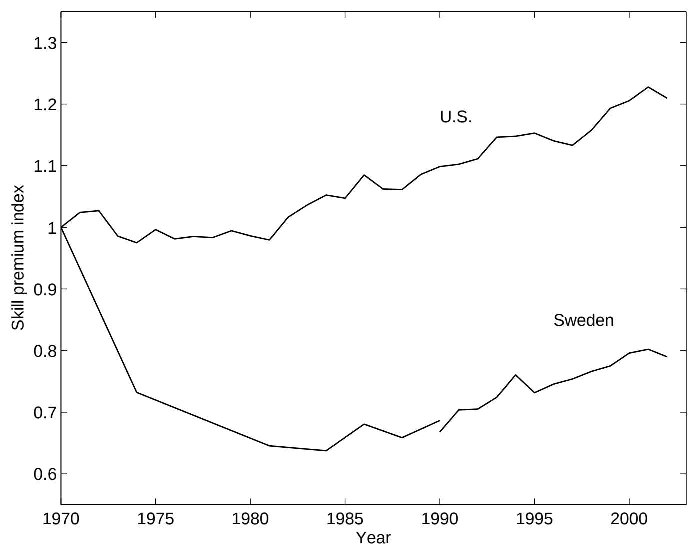

Figure 2: Skill premium in Sweden and the U.S. in the private sector. (The break in the Swedish time series in 1990 reflects subtle changes in the classification of education in the LOUISE database as compared to the census data.)

skill premium remained below its 1970 value, and by 1990, it had fallen by an astonishing 31 percent. Thereafter it recovered somewhat, but in 2002 it was still 20 percent lower than in 1970.

Our task in this paper is to take the labor supply series in Figure 1 as given, and then explain the skill premium series in Figure 2. First, we explore KM's relative wage equation in Section 3 and second, we apply KORV's model with capital-skill complementarity in Section 4.

3 A relative wage equation

Katz and Autor (1999), in their chapter in the Handbook of Labor Economics, review a common approach of studying changes in educational wage differentials by breaking the work force into two broad educational groups, say college equivalents and high school equivalents ('skilled' and 'unskilled' workers). Katz and Autor illustrate the approach by considering a production function with a CES specification over labor inputs, which we here nest inside a Cobb-Douglas specification with capital,

$$y = A\, k^{\alpha} \left[\mu \left(\psi_u h_u\right)^{\sigma} + (1-\mu)\left(\psi_s h_s\right)^{\sigma}\right]^{\frac{1-\alpha}{\sigma}} \quad , \tag{1}$$

where y is aggregate output, k is the stock of capital, and h_u and h_s are hours of unskilled and skilled labor inputs. Time invariant production parameters are α that pins down the capital share of income and σ that determines the elasticity of substitution between skilled and unskilled labor, $1/(1-\sigma)$. The remaining potentially time varying production parameters are as follows. A is a neutral technology factor, μ governs income shares of unskilled and skilled labor, and ψ_u and ψ_s represent unskilled and skilled labor augmenting technology factors.

Under the assumption that skilled and unskilled workers are paid their marginal products, we can use production function (1) to solve for the ratio of marginal products of the two labor types yielding a relationship between relative wages, w_s/w_u, and relative supplies, h_s/h_u, given by

$$\frac{w_s}{w_u} = \frac{1-\mu}{\mu}\left(\frac{\psi_s}{\psi_u}\right)^{\sigma}\left(\frac{h_s}{h_u}\right)^{-(1-\sigma)} \quad . \tag{2}$$

KM provide a successful empirical implementation of this approach to explain changes in the U.S. college/high school wage differential from 1963 to 1987. They assume that a simple linear time trend can approximate the impact of changing parameters $\{\mu, \psi_u, \psi_s\}$ upon the logarithm of the wage differential. After computing the logarithmic change of the skill premium in expression (2) and using KM's time trend representation with their estimated coefficients, we arrive at the following candidate explanation to the skill premium development,

$$\begin{aligned} \log\left(\frac{w_{st}}{w_{ut}}\right) &= \log\left(\frac{1-\mu_t}{\mu_t}\right) + \sigma\log\left(\frac{\psi_{st}}{\psi_{ut}}\right) - (1-\sigma)\log\left(\frac{h_{st}}{h_{ut}}\right) \\ &= 0.033 * \mathit{time} - 0.709\log\left(\frac{h_{st}}{h_{ut}}\right) + \mathit{constant}, \end{aligned} \tag{3}$$

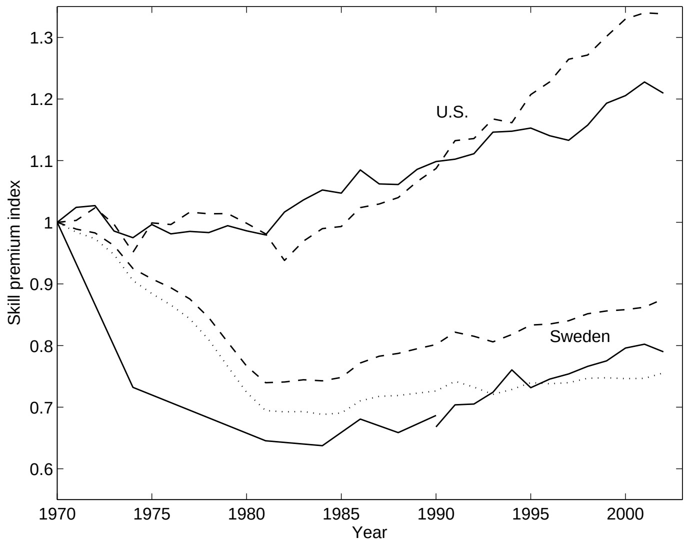

Figure 3: Skill premium in Sweden and the U.S. in the private sector. The solid lines refer to actual skill premia, and the dashed lines depict benchmark simulations of KM's relative wage equation. The dotted line for Sweden is an alternative simulation based on our estimated equation (4).

where KM estimate that the time trend increases the skill premium by 3.3 log points annually and the elasticity of substitution between skilled and unskilled labor is equal to $1/(1-\sigma) = 1/0.709 \approx 1.41$. As noted by Katz and Autor (1999), the estimated elasticity is in the middle of the range of 0.5–2.5 in earlier studies using cross-sectional approaches reviewed by Freeman (1986).

3.1 Model simulation 1970-2002

Figure 3 compares the actual and predicted skill premia for the period 1970-2002 where the predicted skill premia are derived by substituting the ratio of skilled to unskilled workers in the private sector from Figure 1 into KM's equation (3). The model does a very good job in explaining the U.S. skill premia up til the early 1990s, but predicts a more rapid growth in the skill premium than occurred in the last decade.[11]

Concerning the Swedish skill premium, KM's equation (3) predicts a fairly similar time path as that observed in the data, even though the predicted large drop in the 1970s falls short of the even more spectacular observed decline. The

[11]This has previously been noted by Katz and Autor (1999), Card and DiNardo (2002), and Autor et al. (2005), who interpret this as a slowdown in the relative demand for skilled workers. For example, Autor et al. reestimate equation (3) using data from 1963-2003 while allowing for a trend break in 1992. Their results indicate a significant slow-down of demand after 1992, but they also reconfirm the importance of the relative supply growth.

fairly good fit in Figure 3 stands in stark contrast to an earlier attempt by Edin and Holmlund (1995) to estimate equation (3) on Swedish data. They found a much smaller time trend and a much higher elasticity between skilled and unskilled labor of 2.9 which exceeds the range of estimates for other countries, as reported by Katz and Autor (1999) above. The explanation for these disparate estimates is that Edin and Holmlund used a time series of skill premia that turns out to be a poor proxy of actual outcomes; between 1970 and 1990, their skill premium fell only by some 10 percentage points, whereas we showed that the actual skill premium fell by more than 30 percent. When we instead estimate equation (3) using Swedish census data for 1970 and 1990 with interpolations as described above, we obtain estimates remarkably similar to those of KM;[12]

$$\log\left(\frac{w_{st}}{w_{ut}}\right) = \underset{(2.27)}{0.030} * time - \underset{(3.68)}{0.744} \log\left(\frac{h_{st}}{h_{ut}}\right) + constant, \qquad \bar{R}^2 = 0.85, \quad (4)$$

where the numbers in parentheses are absolute t-values. The dotted line in Figure 3 depicts the predicted Swedish skill premium for the period 1970-2002 when substituting the ratio of skilled to unskilled workers in the private sector from Figure 1 into our estimated equation (4).

4 A model with capital-skill complementarity

We now analyze the skill premium in terms of the production function proposed by KORV. They distinguish between capital structures, k_s, and capital equipment, k_e, by augmenting production function (1) as follows,

$$y = A\, k_s^{\alpha} \left\{ \mu \left(\psi_u h_u\right)^{\sigma} + (1-\mu)\left[\lambda k_e^{\rho} + (1-\lambda)\left(\psi_s h_s\right)^{\rho}\right]^{\frac{\sigma}{\rho}} \right\}^{\frac{1-\alpha}{\sigma}}, \qquad (5)$$

where the new parameters λ and ρ determine income shares and substitution possibilities, respectively. The elasticity of substitution between skilled labor and capital equipment is $1/(1-\rho)$, and the elasticity between unskilled labor and equipment or skilled labor is $1/(1-\sigma)$.[13]

Rates of depreciation on structures and equipment are given by δ_s and δ_e, respectively. And as in the standard growth model, KORV assume a linear transformation technology for converting output into new capital but while the transformation rate is constant over time for capital structures, they introduce

[12]Edin and Holmlund (1995) conduct a sensitivity analysis using their estimated wage differential between workers with 16 and 12 years of education which falls almost as much as the actual skill premium, as described in Section 2.2. Though, Edin and Holmlund omit the time trend in that analysis and again, they estimate a very high elasticity of 2.5.

[13]To be precise, the Allen-Uzawa elasticity of substitution (Uzawa, 1962) between unskilled labor and equipment or skilled labor is $1/(1-\sigma)$ and the direct elasticity of substitution (McFadden, 1963) between skilled labor and equipment is $1/(1-\rho)$. The Allen-Uzawa elasticity between skilled labor and equipment is not constant (unless $\sigma = \rho$), nor is the direct elasticity of substitution between unskilled labor and equipment or skilled labor.

equipment-specific technological progress. In the current period, one unit of output can be converted into $1/p_e$ units of capital equipment, which is assumed to increase over time. In a competitive equilibrium, p_e becomes the relative price of capital equipment. Technological progress implies that next period's price p'_e is lower than current period's price p_e.

In contrast to KM, KORV assume that technology parameters $\{\mu, \lambda, \psi_u, \psi_s\}$ are all time invariant, and they note that one of these parameters must be fixed as a normalization in the estimation. To make the latter point explicit, we map the four parameters into three, $\{\Omega, \theta_u, \theta_s\}$, by rewriting production function (5) as follows,

$$y = A\,\Omega k_s^{\alpha} \left\{\theta_u\, h_u^{\sigma} + (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{\sigma}{\rho}}\right\}^{\frac{1-\alpha}{\sigma}} \equiv A\,F(k_s, k_e, h_u, h_s), \tag{6}$$

where $\Omega \equiv (1-\mu)^{\frac{1-\alpha}{\sigma}} \lambda^{\frac{1-\alpha}{\rho}}$, $\theta_u \equiv \mu\psi_u^{\sigma}(1-\mu)^{-1}\lambda^{-\frac{\sigma}{\rho}}$, and $\theta_s \equiv (1-\lambda)\psi_s^{\rho}\lambda^{-1}$.

Given market-determined rental and wage rates, firms' first-order conditions with respect to structures, equipment, unskilled labor and skilled labor can be written as

$$\begin{aligned}
r_s &= A\,\alpha\,F(k_s, k_e, h_u, h_s)\,k_s^{-1}, && (7)\\
r_e &= A\,\Gamma(k_s, k_e, h_u, h_s)\,(k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{\sigma-\rho}{\rho}}\,k_e^{\rho-1}, && (8)\\
w_u &= A\,\Gamma(k_s, k_e, h_u, h_s)\,\theta_u h_u^{\sigma-1}, && (9)\\
w_s &= A\,\Gamma(k_s, k_e, h_u, h_s)\,(k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{\sigma-\rho}{\rho}}\,\theta_s h_s^{\rho-1}, && (10)
\end{aligned}$$

respectively, where

$$\Gamma(k_s, k_e, h_u, h_s) \equiv (1-\alpha)\,\frac{F(k_s, k_e, h_u, h_s)}{\theta_u h_u^{\sigma} + (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{\sigma}{\rho}}}\,.$$

The rental rate r_s on structures and r_e on equipment are such that all capital investments yield the same market-determined rate of return, say a gross rate of return equal to $1+r$, net of physical depreciation and economic obsolescence.[14] Hence, the rental rates satisfy

$$r_s = r + \delta_s, \tag{11}$$

$$\frac{r_e}{p_e} = 1 + r - \frac{p'_e}{p_e}(1-\delta_e) \equiv \tilde{r}_e, \tag{12}$$

where $\tilde{r}_e$ is the marginal product of an amount of equipment that corresponds to the investment of one unit of the good in equipment.

From first-order conditions (9) and (10), we get an equilibrium expression for the skill premium,

$$\frac{w_s}{w_u} = (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{\sigma-\rho}{\rho}}\,\frac{\theta_s}{\theta_u}\,\frac{h_s^{\rho-1}}{h_u^{\sigma-1}}. \tag{13}$$

[14] We follow KORV and ignore risk premia in our analysis of capital investments. Furthermore, to simplify the notation, our expressions are written under perfect foresight.

4.1 Simulation and calibration strategy

As in our simulations based on KM's relative wage equation, we will only use observed time series of the labor skill composition to predict time series of skill premia. Rather than using empirical measures of investments and capital stocks, we will hypothesize that estimates of historical average rates of return determine quantities of capital structures and capital equipment according to firms' first-order conditions. Hence, we first use observed time series of the labor skill composition together with estimates of rental rates to compute implied time series of capital structures and capital equipment from equations (7) and (8). (These implied quantities provide an additional dimension for validation of our calibrated models, as examined in Section 5.3.) Next, given the observed times series of the labor skill composition and the implied capital stocks, we compute predicted skill premia from equation (13).

The simulations are carried out in a calibrated model for each country based on its labor skill composition in 1970. Besides the observed labor skill composition in 1970, our calibration procedure draws upon earlier observations in the literature on the average annual real return (r), depreciation rates (δ_s, δ_e), equipment price movements (p_e'/p_e), and the labor share of output, and estimates of technology parameters (α, σ, ρ). Except for the fact that these averages and estimates reflect economic outcomes over a longer period of time, our calibration of the remaining parameters (Ω, θ_u, θ_s) is guided by the single observation of a country's labor skill composition in 1970. (See Appendix B.2.)

4.2 Parameterization

In the spirit of the calibration approach in quantitative macroeconomics, we draw upon stylized facts and earlier estimates in the literature, especially the estimates of KORV. Our model calibrations for the U.S. and Sweden share the following identical premises, with only one exception being the last premise f) below.

a) The labor share of output is set equal to 2/3, which is a common value in quantitative macroeconomic analysis.

b) The real interest rate is set equal to $r = 0.04$, which is also a standard value in quantitative macroeconomic analysis.

Likewise, KORV estimate the average ex post return on capital structures to 4 percent over the period 1963-1992, but their estimated ex post return series on capital equipment is highly volatile with a mean of 6 percent.

c) Equipment prices have fallen annually by 5 percent over the period 1970-1992, based on a time series constructed by Gordon (1990) until 1983 and by KORV thereafter. Hence, we assume an annual growth rate of equipment prices of $p_e'/p_e = 0.95$.

Gordon (1990) collected detailed information on prices and equipment's characteristics to construct quality-adjusted price indexes, covering the period 1947-1983. On the basis of the historical relationship between Gordon's price indexes and the National Income and Product Accounts (NIPA) official price indexes, KORV extrapolate Gordon's quality-adjusted indexes for 1984-1992. Cummins and Violante (2002) further improve on KORV's extrapolation and extend the series to 2000. Their updated series show that the decline in equipment prices has accelerated over time – the price index fell at an annual rate of about 3 percent until 1975 and reached an annual rate of 6 percent in the 1990s. Our premise above is supported by their estimate that the average annual price decline over the period 1975-2000 was 5 percent. Without any data on quality-adjusted equipment prices in Sweden, it seems reasonable to assume that both countries have faced the same changes in prices, especially since equipment is among the goods that are most traded internationally.

d) Following KORV, annual physical depreciation rates of structures and equipment are set equal to $\delta_s = 0.05$ and $\delta_e = 0.125$, respectively.

The premises in items b)–d) yield rental rates on structures, $r_s = 0.09$, and equipment, $\tilde{r}_e = 0.209$, as given by equations (11) and (12). Since premises b)-d) are assumed to be the same across the U.S. and Sweden, it follows that these rental rates are also the same.

e) We adopt KORV's estimated parameters $\alpha = 0.117$ and $\rho = -0.495$, i.e., capital structures' share of income is 11.7 percent and the substitution elasticity between skilled labor and equipment is 0.67 ($1/(1-\rho)$).

As a sensitivity analysis, we recalibrate the Swedish model below using Swedish data on capital structures' share of income and find that our results are not much affected by the alternative parameterization. Therefore, as a benchmark model, we prefer as far as possible to keep the underlying premises the same across the models for the U.S. and Sweden, to demonstrate that the observed divergence in skill premia is indeed driven by observed labor inputs in the private sector. Though, there is one dimension in which our adoption of a key estimate differs across the models:

f) Concerning the parameter σ, we adopt KM's estimate for Sweden, $\sigma^{\mathrm{KM}} = 0.291$ and KORV's estimate for the U.S., $\sigma^{\mathrm{KORV}} = 0.401$. Hence, the elasticities of substitution between unskilled and skilled labor in Sweden and the U.S. are 1.41 and 1.67, respectively.

Albeit that only the last premise differs across the models for the U.S. and Sweden, it is an important difference to be discussed and motivated below.

Premises a)–f) combined with data on each country's labor skill composition in 1970 allow us to select the remaining parameters $\{\Omega, \theta_u, \theta_s\}$, as described in Appendix B.2.

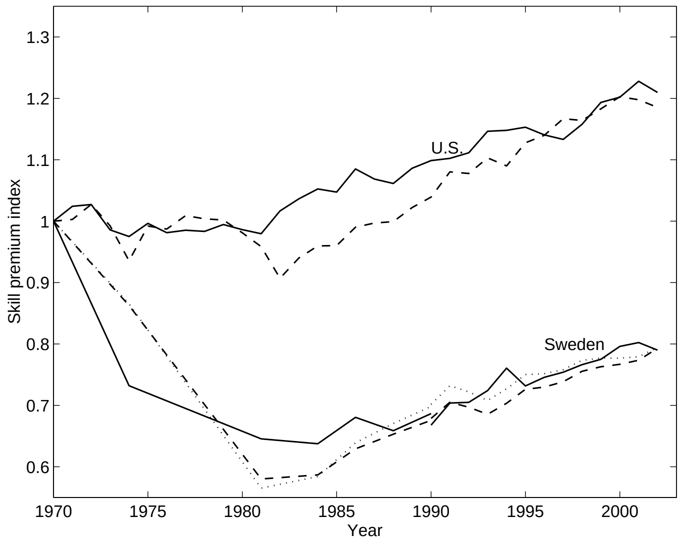

Figure 4: Skill premium in Sweden and the U.S. in the private sector. The solid lines refer to actual skill premia, and the dashed lines depict benchmark simulations in KORV's model. The dotted line for Sweden is an alternative simulation with parameters $\{\alpha, \rho\} = \{.186, -1.25\}$.

4.3 Model simulation 1970-2002

We now use the calibrated models and the observed time series for the labor skill composition over the period 1970-2002 to predict the evolution of skill premia. Figure 4 displays the actual and predicted skill premium evolution in Sweden and the U.S. The predictions are remarkably close to the actual evolution, given that we have calibrated the models using only the labor skill composition in 1970. (Recall that Swedish observations are especially accurate in the years 1970 and 1990-2002 when census data and tax records include the entire population.)

Since our simulations build on the framework of KORV, it is useful to briefly describe how our analysis differs from theirs (besides that we use private sector employment rather than total employment). First, while KORV estimate some key parameters in their study, our analysis is a pure calibration exercise. Based on observations and estimates in the literature, including those of KORV, we calibrate the U.S. and Swedish models to year 1970 which ensures that our models perfectly explain the skill premia in 1970. Second, except for annual observations on the labor skill composition over the period 1970-2002, our simulations do not draw on any other time series such as the data on capital stocks used by KORV. Our simulations can be likened to balanced-growth paths perturbed by variations in the labor skill composition. Third, the balanced-growth part of our simulations is driven by the premises of a constant rate of decline in equipment prices and a constant rate of return on capital. In addition to skill premia, our simulations predict stocks of structures and equipment. In contrast to KORV who use capital stocks as inputs in the simulations, we will compare the implied investment rates from our simulations to data in order to assess the performance of our models.

5 Sensitivity analysis and other implications

5.1 Sensitivity with respect to elasticity of substitution

The skill premium evolution in Sweden and the U.S. is surprisingly well explained by a common framework that shares five of our six premises underlying the parameterization in section 4.2. The only difference, premise f), concerns the elasticity of substitution between unskilled and skilled labor. In particular, our model for Sweden adopts KM's estimated elasticity of 1.41, while our model for the U.S. uses KORV's estimated elasticity of 1.67. This difference in elasticities might seem small relative to the prevailing range of estimates in the empirical literature. As noted above, Katz and Autor (1999) report that cross-sectional and time series studies lend support to a range of 0.5-2.5. However, in the parameter space of our structural framework, the difference between an elasticity of 1.41 versus 1.67 is large.

To provide a metric for the significance of the postulated difference in the elasticity of substitution between unskilled and skilled labor in Sweden versus the U.S., we explore the implications of varying that elasticity. Specifically, keeping the same premises as in the benchmark models except for varying the parameter σ, we calibrate a continuum of new models for Sweden and the U.S., respectively. These models are then simulated to predict the skill premium in 1990 and 2002, as depicted in Figure 5 by the dashed lines for Sweden and the solid lines for the U.S. The simulations of the benchmark models for Sweden and the U.S. are identified by circles. The simulated skill premia should be compared to the actual outcomes in 1990 and 2002, as represented by the dotted lines (and also marked with crosses at the benchmark elasticities of substitution). The distance between a circle and a cross for a given country and year is the same as the corresponding difference between the actual and simulated skill premium for that country and year in Figure 4. As we already know, the benchmark model for Sweden with an elasticity of 1.41 almost perfectly explains the Swedish skill premium in 1990 and 2002. But if we instead had assumed an elasticity of 1.67, the model for Sweden would have erroneously predicted that the Swedish skill premium in 1990 had reverted to its 1970 level, followed by a further explosive increase in 2002. Similarly, the benchmark model for the U.S. with an elasticity of 1.67 explains fairly well the actual skill premium evolution in the U.S. But if we instead had assumed an elasticity of 1.41, the model for the U.S. would have counterfactually predicted a depressed U.S. skill premium in both 1990 and 2002 relative to its 1970 level.

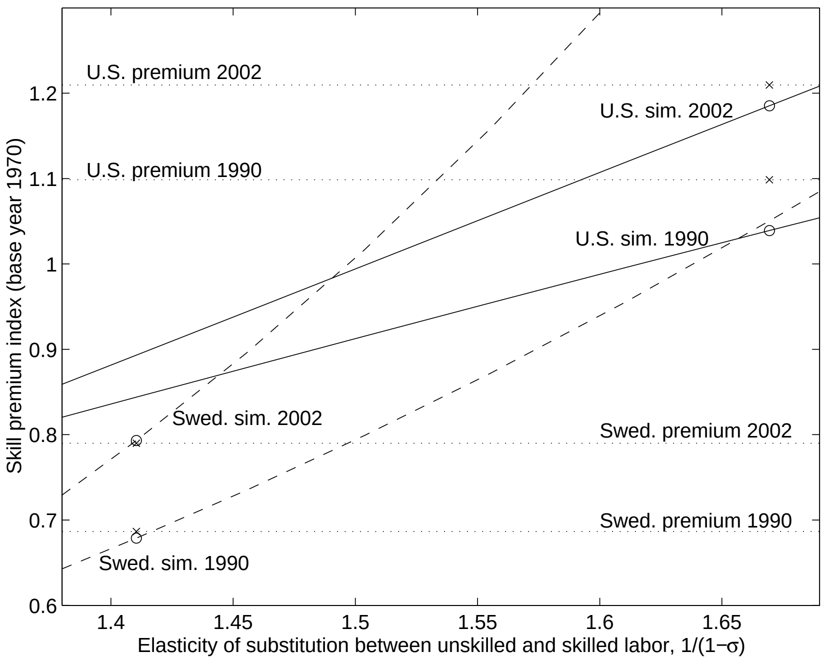

Figure 5: Sensitivity analysis with respect to the elasticity of substitution between unskilled and skilled labor, $1/(1-\sigma)$. Keeping the same premises as in the benchmark models except for varying the parameter σ, we calibrate a continuum of models for Sweden and the U.S., respectively. These models are then simulated to predict the skill premium in 1990 and 2002, as depicted by the dashed lines for Sweden and the solid lines for the U.S. The simulations of the benchmark models for Sweden and the U.S. are identified by circles. (Recall that the elasticities of substitution in Sweden and the U.S. are assumed to be 1.41 and 1.67, respectively.) The actual skill premia in 1990 and 2002 are represented by the dotted lines (and also marked with crosses at the benchmark elasticities of substitution).

We conclude that production function (5) and the time series of labor inputs can only explain observed skill premia in Sweden and the U.S. if we allow for differences in the elasticity of substitution. On the empirical side, Bergström and Panas (1992) provide evidence for a relatively low elasticity of substitution between unskilled and skilled labor in Swedish industries.[15] On the theoretical side, we conjecture that the much lower absolute fraction of skilled labor in the labor force in Sweden as compared to the U.S., indicates differences in education systems which manifest as a lower elasticity of substitution in our parsimonious representation of the production function.[16] If that conjecture is correct, the fewer 'skilled' workers in Sweden represent an input that is less substitutable for 'unskilled' workers as compared to the corresponding categories in the U.S.

5.2 Sensitivity with respect to capital share parameter

To demonstrate robustness of our analysis in one dimension of the parameter space, we consider the fact that the national income accounts suggest a higher capital structures' share of output in Sweden, $\alpha = .186$ in the private sector (which is, incidentally, similar to the share of 0.18 in the aggregate economy).[17] We now adopt this alternative parameter value $\alpha = 0.186$ but also adjust the elasticity of substitution between skilled labor and capital equipment, $1/(1-\rho)$, to accommodate this different technology specification. In particular, we calibrate the value $\rho = -1.25$ that allows us to retain the good match with the actual Swedish skill premium in 2002. It should be noted that this value is well within the range of the empirical estimates (see Hamermesh, 1993).

The dotted line in Figure 4 depicts this alternative parameterization of the model for Sweden, and we see that it explains the evolution of the Swedish skill premium as well as the benchmark model. But while this latter recalibration utilizes information about actual outcomes in both 1970 and 2002, the benchmark parameterization is only based on the labor skill composition in 1970.

5.3 Implications for labor share, investment and growth

The model has several other implications, as reported in Table 2. The first implication concerns the labor share of output. As noted by KORV, the specification of the production function does not guarantee a constant labor share of output. But the model for the U.S. yields a labor share that remains at roughly

[15]Bergström and Panas (1992) estimate separate elasticities for 4 years (1963, 1968, 1974 and 1980) and for 8 manufacturing industries in Sweden. The estimates are fairly stable across time for each industry and the average elasticity in each of the 8 industries are 0.36, 0.47, 0.91, 1.25, 1.26, 1.31, 1.45, and 3.50.

[16]Using our weighting scheme for computing quantities of labor inputs, the skilled fraction of the labor force in the private sector in 1970 (2002) constituted 3.6% (16.4%) in Sweden as compared to 16.4% (36.6%) in the U.S.

[17]See Statistics Sweden's web site www.scb.se for data after 1994, while earlier data is published in Statistical reports N1981:2:5 (Appendix 2, "Capital formation and stocks of fixed capital 1963-1980"; Appendix 4, "Production and factor income 1963-1980") and N 10 SM 9501 ("National accounts 1980-1994").

two-thirds of output throughout the period. For Sweden on the other hand the model predicts a substantial fall in the labor share of output, whereas it remains fairly stable in the data, though somewhat lower in the 1990s compared to the 1970s (see footnote 17 for sources). While a serious flaw, it can easily be corrected. Under our alternative parameterization of Section 5.2, the fall in the labor share mirrors that in the data and at the same time the evolution of the skill premium predicted by the model still closely conforms with that observed in the data.

Table 2: Model simulations

	U.S.	Sweden	
	benchmark	benchmark	alternative
Labor share of output in 2002 (equal to 0.667 in 1970)	0.645	0.556	0.630
Annual growth rate of output[a] (and of structures)	2.0%	2.9%	1.6%
Annual growth rate of equipment[a]	7.7%	9.8%	7.7%

[a] Growth rates 'per worker,' i.e., we normalize $h_{ut} + h_{st} = 1$.

The second implication concerns productivity growth. All productivity gains in our analysis can be explained by equipment-specific technological change, as earlier shown by Greenwood et al. (1997).[18] In the model for the U.S., output per worker grows by 2 percent per year, which is in line with what is observed for the U.S. For Sweden, the model predicts higher productivity growth, whereas actual growth has been a bit smaller.[19] Again this can be corrected using the alternative parameterization.

The third implication concerns the growth rate of equipment. The model for the U.S. predicts that the stock of equipment in efficiency units grow at an annual rate of 7.7 percent which is very similar to the 7.4 percent reported in KORV based on Gordon (1990) capital series. The model for Sweden suggests that Swedish equipment grew at least as fast. How does this fit the facts? While there exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon, Statistics Sweden reports data on capital stocks measured using similar methods as in the U.S. NIPA. KORV reports that U.S. equipment, as measured in NIPA, grew at an annual rate of 3.4 percent between 1980-92. According to Statistics Sweden, the stock of equipment grew at 3.5 percent per year during the same period.[20] While not conclusive, this suggests that the

[18] According to Figure 3 in Greenwood et al. (1997), neutral technological change has on average contributed negatively to productivity growth since the 1970s and hence, equipment-specific technological change has become the source of all productivity gains.

[19] Using data from Penn World Table Version 6.1, the average growth rate in Real GDP per worker was 1.7 percent and 1.2 percent in the United states and in Sweden, respectively.

[20] See Statistical report N 10 SM 9501 ("National Accounts 1980-1994": Appendix 3, "Stocks of fixed assets and national wealth 1980-1995").

stocks of equipment developed similarly in both countries and once again, the alternative parameterization of Sweden looks compelling.

5.4 Earlier study using the same framework

The framework of KORV has recently been applied on Swedish data by Lindquist (2005), though he offers a starkly different view on how the Swedish skill premium has evolved during the last three decades. When estimating production function (5), Lindquist uses a poor proxy of the skill premium which exhibits too little compression; in 1999 his measure is back to its 1970 value whereas we showed, using census data and tax records, that it is some 20 percent below. It is therefore surprising to note that his estimates for key parameters such as $\sigma = 0.29$ and $\rho = -0.93$ are very similar to those used in this paper. The most likely explanation is that he uses the total labor force rather than private sector employment when computing the labor skill composition. Hence, Lindquist's severe underestimate of the true fall in the Swedish skill premium is offset with another large underestimate of the increase in the skilled to unskilled labor ratio that is relevant for wage determination. As we have argued, it is the withdrawal of low-skilled workers from the private sector into public sector employment that can rationalize the true dramatic decline in the Swedish skill premium.

6 Discussion

6.1 Public sector employment

We are not the first ones to suggest that public sector employment can be an important determinant for wage outcomes.[21] But we are not aware of any other quantitative study that uses a structural model and time series data to seek empirical support for such a hypothesis. Our findings show that public sector employment can indeed explain why the Swedish skill premium has taken such a different path as compared to that of the U.S. since 1970.

In his account of the Swedish employment experience, Lindbeck (1997, pp. 1280, 1311) argues that "the dramatic expansion of public-sector employment is partly a consequence of the government serving as 'an employer of last resort.' ... the expansion of permanent *public-sector employment* by about 600,000 people (15 percent of the labor force) from 1970 to 1985 ... kept up the demand for low and medium-skilled" workers. Figure 6 shows that the public-sector share of total labor input has virtually exploded in Sweden while the corresponding share in the U.S. is rather flat. Though, as we have argued above and as implied by Lindbeck's remarks, the expansion of the public sector gains significance through its impact on the relative supplies of unskilled and skilled labor in the private

[21]Edin and Topel, whom we quoted in the Introduction to attribute the Swedish wage compression to a collusion between the trade union and the employers' association, also note the potential importance of public sector growth. "The government may have supported the price of low-wage workers by simply hiring them, soaking up the excess of supply over private demand" (Edin and Topel, 1997, p. 190).

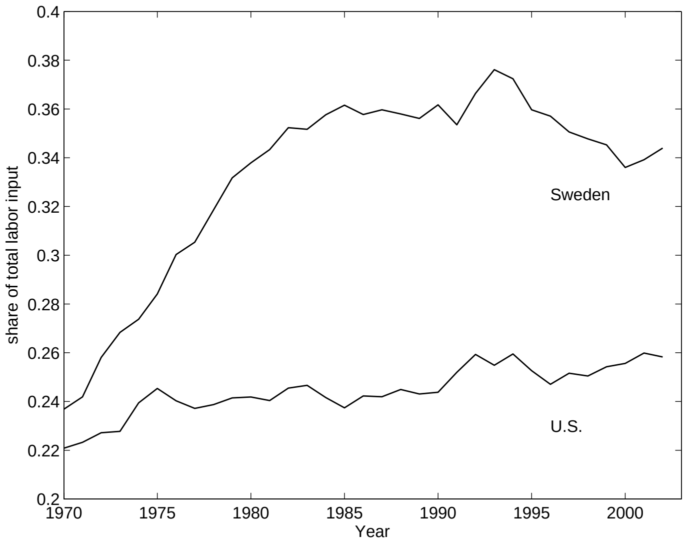

Figure 6: Public employment share in Sweden and the U.S.

sector. Figure 7 shows how the disproportionate hiring of unskilled workers in the Swedish public sector contributed to a sharply increasing private-sector ratio of skilled to unskilled labor in Figure 1. Specifically, Figure 7 depicts the ratio of skilled/unskilled labor in the private sector relative to that in total employment. While the U.S. have seen a modest increase in this relationship, the ratio of skilled to unskilled labor in the Swedish private sector has evidently increased sharply because of a disproportionate hiring of unskilled workers in the public sector.

As noted by Freeman et al. (1997, p. 13), essentially all Swedish employment growth in the 1970s and 1980s was in services provided by the local government. And one large component was publicly provided day care for preschool children that "has grown explosively since the mid-1970s to be almost half as large as the employment in the education sector." In the words of Rosen (1997, pp. 81, 80), "Sweden has 'monetized' the household sector of its economy by substituting publicly for privately produced household services on a grand scale in the past three decades. ... the welfare state encourages extra production of household goods and discourages production of material goods. From the normative view of economic efficiency, too many people provide paid household (family) services for other people, and too few are employed in the production of material goods. From the view of positive economic analysis, this is what explains the growth of local government employment." Our analysis shows how the forces identified by Rosen have also had equilibrium repercussions on the Swedish skill premium.

To illustrate the difficulty of explaining the Swedish skill premium without removing public sector employment, Figure 8 depicts the predictions of KORV's

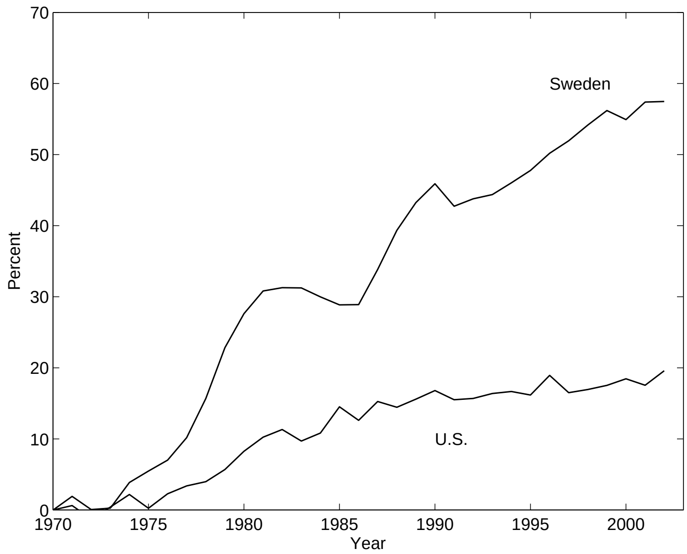

Figure 7: Ratio of skilled/unskilled in private sector relative to ratio of skilled/unskilled in total employment (percentage change since 1970).

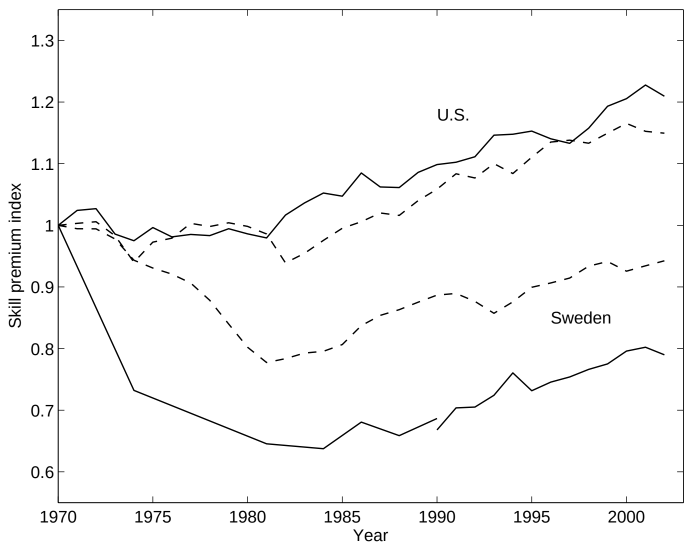

Figure 8: Using total employment to predict skill premia with KORV's model. The solid lines refer to actual skill premia, and the dashed lines depict predicted skill premia using the benchmark calibration procedure.

model when using total employment and the benchmark calibration procedure. As we have indicated earlier, it makes hardly any difference for the U.S. whether private or total employment is used but total employment in Sweden cannot explain the dramatic decline of the Swedish skill premium.[22] The model predicts a Swedish skill premium in 2002 just short of its 1970 value, and not 20 percent below as in the data.[23] The results are even more extreme when using total employment in KM's relative wage equation. It counterfactually predicts that the Swedish skill premia should have risen by 20 percent in 2002 relative to 1970.

6.2 Market versus nonmarket forces

Our hypothesis that the expansion of the Swedish public sector can explain the dramatic decline of the skill premium between 1970 and 1990 is of course a theory about nonmarket outcomes. Tax-financed local governments have disproportionately hired unskilled workers in large numbers and thereby affected the ratio of skilled to unskilled labor available to the private sector. But in contrast to other nonmarket theories that imply wages inconsistent with market clearing, both unskilled and skilled labor in our analysis are paid their marginal products in the private sector. Hence, given the quantities of labor available in the private sector, workers receive the same compensation as they would in a competitive equilibrium.

There are two ways that the allocation of labor between the public and the private sector could have arisen, and our analysis does not discriminate between them. First, suppose that trade unions and employer federations have set the observed wages. As long as the public sector stands ready to hire any surplus of workers at those wages (by acting as 'an employer of last resort'), the observed labor quantities in the private sector are consistent with our theory, i.e., private

[22]While Figure 8 contains the actual skill premium in the private sector rather than that for total employment, the two time series for the U.S. are virtually identical. In Sweden, the skill premium for total employment does gradually fall behind the private skill premium starting around 1990 with an eventual gap of 7 percentage points in 2002 as reported in Table 1 and hence, the former time series would exacerbate the prediction error in Figure 8.

[23]We have experimented with several different calibration procedures to investigate if a reasonable parameterization exists for which predictions based on total employment will match the evolution of the Swedish skill premium, without having counterfactual predictions along other dimensions. We have not been able to find such a parameterization. First, under the benchmark calibration procedure we allowed σ to vary between 0.2 and 0.4. Note that a higher σ makes the skill premium fall less during the 1970s and 1980s and increase more during the 1990s. For any σ less than 0.205 the predicted skill premium in 1990 is too low. For any σ larger than .268 the skill premium is too large in 2002. For any intermediate value of σ the skill premium is counterfactually predicted to rise rapidly between 1985 and 1990 and then remain approximately constant during the 1990s. The reason is that the relative supply of skilled to unskilled labor did not increase between 1985 and 1990 in aggregate employment whereas it did in private employment, as implied by Figures 1 and 7. Second, we repeated the same exercise but with $\alpha = 0.186$ as suggested in our alternative calibration. The results were unchanged. Third, with $\alpha = 0.186$, we devised flexible versions of our alternative calibration procedure in which σ and ρ were varied to attempt to match the evolution of both the skill premium and the labor share, but to no vail.

employment is such that marginal products of unskilled and skilled labor are equal to the pre-determined wages. Second, suppose instead that the political process determines the size of the public sector and its hiring of workers. A well-functioning market economy would then ensure that the remaining labor quantities available to the private sector are fully employed at market-clearing wages equal to the marginal products of unskilled and skilled labor.

Our theory cannot discriminate between these alternative scenarios because, like KM and KORV, we take the quantities of unskilled and skilled labor as given. Moreover, like those earlier studies, it is a shortcoming that we do not endogenize the economies' skill formation and how such skill formation responds to changing rates of return to education as implied by a changing skill premium.[24] Though, as an explanation of the observed skill premium evolution in Sweden and the U.S., we conclude that observed quantities of unskilled and skilled labor in the private sector in each country are consistent with labor being paid its marginal product. This is a startling conclusion given the dramatic decline in the Swedish skill premium until the mid-1980s, as we have documented with census data and tax records for the economy's entire population.

6.3 Implications for Europe

While our theory attributes the decline in the Swedish skill premium to an expanding public sector that has disproportionately hired unskilled workers, other ways of withdrawing unskilled labor from the private sector would also reduce the skill premium. Such alternative ways might include the increasing number of European workers on disability and unemployment insurance, and in early retirement programs. OECD (2003, Table 4.1) reports large increases in the proportion of working-age people receiving income-replacement benefits in Europe between 1980 and 1999. For example, that benefit dependency rate in full-time equivalents rose from 14% and 15% in France and Germany to 24% and 22% percent, respectively. Meanwhile, the benefit dependency rate in the U.S. has slightly decreased over these two decades to less than 14% of the working-age population in 1999.

There are two reasons for suspecting that unskilled labor is overrepresented in the increasing number of benefit recipients in Europe. First, the cost to individual workers of enlisting in welfare programs instead of working depends on the fraction of labor market earnings that benefits are supposed to replace. Since most income-replacement programs have some form of ceilings on benefit levels, unskilled labor with their lower income levels typically face higher effective replacement rates than those of skilled labor. Second, the opportunity cost of withdrawing from labor market participation does not only depend on individual workers' current wages but also anticipations about their future

[24] It is clearly desirable to explicitly model the supply of skilled workers but we acknowledge that this might be especially complicated in the case of Sweden where both the costs of and returns to education have been 'socialized' in form of free education with generous general student grants and a labor income tax system that has been highly progressive in the 1970s and 1980s.

income growth. There is evidence suggesting that the earnings prospects of workers at the lower end of the skill spectrum have been subject to downward pressures that have been attributed to various causes including globaliziation and skill-biased technological change.

As an example of both these reasons operating in a U.S. context, Autor and Duggan (2003) argue that reduced screening stringency and rising replacement rates of the disability insurance program in the U.S., have led to an increasing number of unskilled workers exiting the labor force to seek disability benefits in recent decades of adverse demand shifts for their labor services. Because of the progressive (i.e., concave) benefit formula, they find that it is primarily high school dropouts who collect disability insurance in the U.S. In a European context with a more expansive welfare state, we hypothesize that Autor and Duggan's characterization of U.S. disability insurance and its impact on unskilled labor represents a 'microcosmos' of what has occurred in Europe in the last three decades. If that is so, large withdrawals of unskilled labor into European welfare programs and the resulting relative scarcity of unskilled labor in the private sector might explain why skill premia have risen so much less in Europe than in the U.S.[25] A hypothesis that merits further investigation.

[25]Our hypothesis is related to a commonly voiced argument that European productivity gains might partly be "artificial, as Europe made labor expensive ... firms were forced to slide northwest up their labor demand curves, retaining high productivity workers while forcing low-productivity workers into unemployment or out of the labor force entirely" (Gordon, 2006). Both hypotheses emphasize outcomes where unskilled or low-productivity workers in Europe are induced to leave the ranks of the employed. Another theory consistent with this view is Ljungqvist and Sargent's (1998) supply explanation to high European nonemployment which draws upon evidence of increased turbulence for individual workers. Their model of "ex ante identical individuals who can accumulate human capital only through work experience is best thought of as a model of blue-collar workers. ... The fact that welfare benefits are based on past earnings causes [workers with disadvantageous labor market outcomes] literally to 'bail out' from the active labor force" (p. 547).

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Appendix A Data construction

This appendix describes how we construct our measures for skill premia and labor inputs. The source for our U.S. data are the 1971-2003 March CPS Annual Demographic Survey files provided by UNICON, from which we extract data for the years 1970-2002. The source for our Swedish data are the Census of Population Surveys for 1970 and 1990, and the LOUISE database for the period 1990-2002. The Census of Population Surveys cover *all* Swedish individuals, and the LOUISE database covers all individuals between 16 and 64 years of age for the period 1990-95 and all individuals above age 16 thereafter.

We construct our measures for skill premia and labor inputs using two different samples. The supply sample includes all workers between 16 and 70 years of age. The wage sample is restricted to full-time workers. In each sample, we sort workers into groups according to age (five-year intervals), sex, race (only U.S.; white, black and other), education (no high school, high school diploma, some college and college graduate), and industry classification of their main employment. We use the industry variable to group individuals into (i) private sector employment (manufacturing and services), and (ii) public sector employment (health, education, postal services and government administration). While health and to some extent education are largely produced in the private sector in the U.S., they are almost exclusively produced in the public sector in Sweden. To facilitate comparison between the two countries we choose to use the Swedish division into private and public employment for both countries. Note however that for the U.S., the distinction between private and public employment is immaterial for the results.

Using the wage sample we calculate with-in group average hourly wages as the ratio between total income and total annual hours for each group. Using the supply sample we calculate total annual hours for each group. The groups are then sorted into two classes; skilled and unskilled labor, where by skilled we mean college graduates. We obtain class-specific averages by aggregating across groups. For the aggregation we assume that groups are perfect substitutes within a class and use average group wages as weights.

All variables used are listed in Appendix A.6.

Appendix A.1 U.S. data

We first construct the supply sample by excluding (i) individuals below age 16 and above age 70, (ii) individuals who are not in the labor force, (iii) individuals who work without pay and (iv) individuals without education or industry classification. To give an indication of the sample size, the number of remaining observations are 51514 in 1970, 74851 in 1980, 67732 in 1990 and 94371 in 2002.

In the wage sample we follow Autor et al. (2005) by excluding (i) the self-employed, (ii) individuals with less than 40 weeks worked per year, (iii) individuals who work less than 35 hours per week, (iv) individuals with allocated income, (v) individuals whose weekly pay (after topcode adjustment, see below) is less than \$67 in 1982 dollars (using CPI as the deflator), or whose hourly wage exceeds 1/35th the topcoded value of weekly earnings.[26] The number of remaining observations are 28308 in 1970, 42433 in 1980, 48664 in 1990 and 70319 in 2000.

We use the following classification scheme for education. We divide individuals into 4 groups; (i) less than 12 years of schooling, (ii) high-school graduate (completed 12 years of schooling), (iii) some college, and (iv) college graduate. College graduate include individuals who have completed 16 years of schooling. In 1992 there was a change in the recording of educational attainment. To keep consistency in classification we follow the suggestion in the UNICON documentation and classify individuals whose 13th year of schooling is not completed as 'with some college'.

We classify individuals as private sector employees if their CPS industry classification belongs to the following set of CPS codes; (17-817, 849, and 888-899) for the years 1970-81, (10-411, 420-811, 841, and 882-893) for the years 1982-2001 and (111-6290, 6380-7790, 8560-9090, 9190 and 9290) for the year 2002.

Prior to March 1989, wage and salary income is reported in the CPS as a single variable that is topcoded at values between $50,000 and $99,999. Beginning in 1989, wage and salary income is reported in two separate variables, corresponding to primary and secondary earnings, which are topcoded separately.[27] Like Autor et al. (2005), we handle topcoded earners as follows. We adjust primary and secondary earnings separately before summing them. For all years we multiply the topcoded values by 1.5. See Appendix A.4 for a sensitivity analysis with respect to different procedures of handling topcoded earners.

Appendix A.2 Swedish data

We construct the supply sample by excluding (i) individuals below age 16 and above age 64, (ii) individuals who are not in the labor force, and (iii) individuals without education or industry classification. The number of remaining observations are approximately 3.2 million in 1970, and between 3.6 and 3.9 million each year 1990-2002. We construct the wage sample by excluding the self-employed.

We classify educational attainment as in the U.S. data with two exceptions. First, a high-school diploma can be obtained in Sweden after 11 or 12 years of schooling and we regard both types of graduates as high school graduates.

[26] We have investigated the following alternative specifications with unchanged results; (i) excluding individuals whose weekly earnings are less than 40 times half the hourly minimum wage, (ii) including individuals whose weekly earnings are less than 40 times half the hourly minimum wage but imputing their wages to equal half the minimum wage, or (iii) excluding individuals whose weekly earnings are less than 40 times one sixth of the hourly minimum wage.

[27] Between 1989 and 1995 both primary and secondary earnings were topcoded at $99,999. Between 1996 and 2002 primary and secondary earnings were topcoded at $150,000 and $25,000, respectively. In 2003 these values were raised to $200,000 and $35,000, respectively. Beginning in 1996, topcoded earners are assigned the mean of topcoded earners in the same sociodemographic group (sorted according to sex, race and worker status). We follow Autor et al. (2005) and reassign primary and secondary earnings to their topcoded values.

Second, obtaining a traditional Swedish college degree requires 3-5 years of studies depending on field of specialization. We thus classify individuals who have completed a 3-year university or college degree as college graduates.

Finally, for private sector employment we use SNI69 codes (10-71, 81-83, 92 and 94-95) in the Census data for 1970 and 1990, SNI92 codes (0-64110, 64204-73000, 74111-75000, and 92000-95001) in the LOUISE database for 1990-2001, and SNI2002 codes (0-64110, 64204-73000, 74111-75000, and 92000-95001) in the LOUISE database for 2002.

Appendix A.3 Skill premia and labor supplies

For each individual i and year t, we use observations on the number of hours worked per week, n_{it}^{hrs}, the number of weeks worked per year, n_{it}^{wks}, annual earnings, e_{it}, and the sampling weight υ_{it}. Based on individual characteristics, the workers are divided into groups $g \in G$. Under our assumption that workers are perfect substitutes within a group, we compute the hourly wage rate W_{gt} for group g in year t as

$$W_{gt} = \frac{\sum_{i \in g} e_{it} \upsilon_{it} d_{it}}{\sum_{i \in g} n_{it}^{wks} n_{it}^{hrs} \upsilon_{it} d_{it}} ,$$

where $d_{it} = 1$ if individual i in year t belongs to the wage sample, and zero otherwise. Next, we calculate the average annual hours worked H_{gt} for group g in year t as

$$H_{gt} = \frac{\sum_{i \in g} n_{it}^{wks} n_{it}^{hrs} \upsilon_{it}}{\upsilon_{gt}} ,$$

where $\upsilon_{gt} = \sum_{i \in g} \upsilon_{it}$.

The groups are then aggregated into two classes; 'skilled' workers, G_s, and 'unskilled' workers, G_u. For each class $j \in \{s, u\}$ and year t, the class labor supply is computed as $h_{jt} = \sum_{g \in G_j} H_{gt} \bar{W}_g \upsilon_{gt}$, where the weights $\bar{W}_g$ are the average group wages for 1970–1990;[28] and the class wage rate is calculated as $w_{jt} = \sum_{g \in G_j} H_{gt} W_{gt} \upsilon_{gt} \; / \; h_{jt}$.

Several remarks follow. First, in the U.S. data we use the CPS sampling weights but since the Swedish data includes all individuals we simply count heads ($\upsilon_{it} = 1$). Second, in the U.S. data beginning in 1976 'hours worked' refers to 'usual hours worked' but prior to 1976 we only have data on 'hours worked last week.' Prior to 1976 some individuals in the supply sample may thus have worked last year but were unemployed or not at work last week. Since we assume that workers are perfect substitutes within a group, the hours worked

[28] As a sensitivity analysis, we have tried three alternative weighting schemes; 1970 group wages, 1990 group wages, and 2002 group wages, respectively. The results were similar.

for these individuals are estimated as

$$n_{it}^{hrs} = \frac{e_{it}}{W_{gt}\, n_{it}^{wks}}, \qquad \text{for } i \in g \text{ and } d_{it} = 0.$$

Third, in the U.S. data prior to 1976 the number of weeks worked were reported in intervals (0, 1-13, 14-26 weeks etc.) while beginning in 1976 it was reported in actual weeks. Prior to 1976 we have assigned an exact number of weeks using with-in interval averages based on CPS data from 1976-1980.

Fourth, in the Swedish data for 1970 and 1990 hours worked is reported in intervals. For 1970 the intervals are 1-19, 20-34 and ≥35 hours per week, and for 1990 the intervals are 1-15, 16-19, 20-34, ≥35. We have assigned hours per week using midpoints, assuming 43 hours per week for 1970 and 40 hours per week for 1990 in the upper interval.[29] A potential problem with this procedure is that the implied average numbers of hours per week do not exactly match the ones reported by Statistics Sweden (see www.scb.se; and Pohjolassa, 1983). In the latter data, the average hours worked in 1970 (1990) was 39.3 (36.8) hours per week, with men working 43.5 (40.4) and women 32.8 (32.7) hours per week. Using midpoints implies that average hours worked in 1970 (1990) was 39.9 (36.4) hours per week, with men working 42.6 (38.7) and women 35.8 (33.9) hours per week. That is, using midpoints overestimates the working hours of women and underestimates the working hours of men. To investigate the importance of this mismatch we perform the following robustness check. The midpoints in each interval is multiplied by two factors; the first capturing aggregate time affects and second capturing time specific gender affects. We calibrate these factors so that average hours worked by men, by women and averaged across gender in both years in our data match average hours work per week as reported by Statistics Sweden. The results of the paper were not affected.

Finally, in the Swedish data for the years 1990-2002 we only have observations for hours worked in 1990. To obtain hours for 1991-2002 we assume that average hours work in each group g has remained constant at its 1990 level. This is roughly consistent with hours worked by age and gender as reported by Statistics Sweden (see www.scb.se). Between 1990 and 2000 average hours worked for men remained constant while it increased by 3 percent for women. In the age dimension, only a small fraction of the labor force experienced large changes; individuals age 19 and younger.

Appendix A.4 Topcoding in U.S. data

As mentioned above, data on wage and salary income is topcoded in U.S. data. Figure 9 shows how the share of individuals who are topcoded in our wage sample has increased over time (but whenever the topcoded value has been raised, there is a sharp drop). We therefore believe it is important to investigate the

[29] In the upper interval, we choose these hours to match the maximum number of regular working hours per week stipulated by law. While the 40 hour week was introduced in 1970, it was not implemented until 1973.

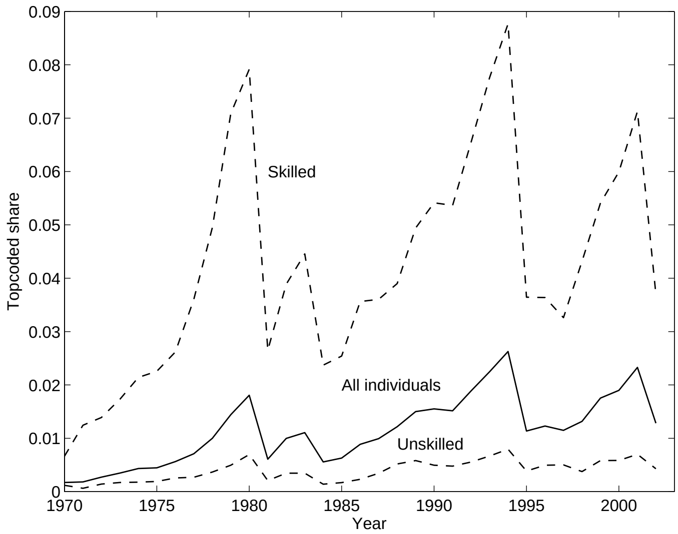

Figure 9: Share of individuals who are topcoded. The solid line depicts the topcoded share among all individuals while the upper (lower) dashed line shows the topcoded share among skilled (unskilled) labor. The topcode classification is based on total earnings until 1988, and primary earnings afterwards.

sensitivity of the U.S. skill premia to different procedures of handling topcoded earners. Besides our benchmark procedure of multiplying topcoded values by the adjustment factor 1.5, we consider three alternative procedures. In the first procedure, denoted NO, we follow KORV and keep the topcoded values without multiplying by any adjustment factor. In the second procedure, denoted CD, we follow Card and DiNardo (2002) who recensor primary and secondary earnings after 1989 to the 1988 topcoded values of \$99,999 and \$25,000, respectively, and then like KORV use topcoded values without multiplying by any adjustment factor. In the third procedure, denoted P, we assume that the top decile of the earnings distribution is Pareto distributed,[30] and multiply topcoded values by an estimated adjustment factor that is implied by the Pareto distribution.

Under the Pareto assumption, the cumulative distribution function is given by $F(e) = 1 - (b/e)^a$, which implies that the ratio between the average income above e and e is equal to $a/(a-1)$. Hence, knowledge of the shape parameter a is sufficient for constructing an adjustment factor for our topcoded earnings. We obtain year-spefic estimates of a_t as follows. Let $\vartheta_t(e)$ denote the fraction of earners with income greater than e in year t, i.e., $\vartheta_t(e) = (b_t/e)^{a_t}$, and after taking logarithms, $\log \vartheta_t(e) = constant_t - a_t \log e$. Based on this relationship between the fraction of earners with income above e and e, we compute OLS

[30]For a recent study using the Pareto distribution to interpolate income distributions, see Piketty and Saez (2003).

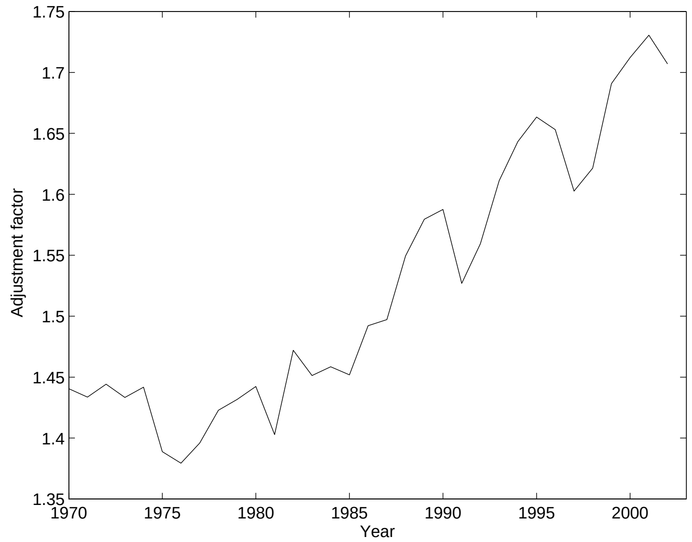

Figure 10: The adjustment factor $a_t/(a_t - 1)$ that topcoded earnings are mulitiplied by under the assumption that the top decile of the earnings distribution is Pareto distributed.

estimates of a_t. The data points $\{\vartheta_t(e_{it}), e_{it}\}$ are constructed from the earners in the top decile who are not topcoded. (The R^2 values in these regressions are above 0.98.) Figure 10 shows the implied annual estimates of the adjustment factor, $a_t/(a_t - 1)$, which are then used in our Pareto topcoding procedure. The adjustment factor is approximately 1.45 until 1985 and then gradually increases towards 1.7.

Figure 11 shows the U.S. skill premium for the benchmark and the three alternative procedures. The benchmark and the Pareto procedure yield very similar results. The NO procedure produces a skill premia that is always below the benchmark and the Pareto procedure. This is due to the fact that there are relatively more skilled individuals that are topcoded as shown in Figure 9. This is also why the CD procedure predicts that the increase in the skill premium was an episodic event of the 1980s without much change in the 1990s; there are relatively more skilled individuals that have their earnings recensored to their 1988 values.

Appendix A.5 Swedish labor supply between 1971-1989

To construct Swedish labor inputs for the years 1971 through 1989 we use data from labor-market surveys (*Arbetskraftsundersökning*, AKU). Individuals with SNI codes 91,931-934 are classified as public sector employees. Given our Census-based measures of skilled and unskilled labor inputs in 1970 and 1990,

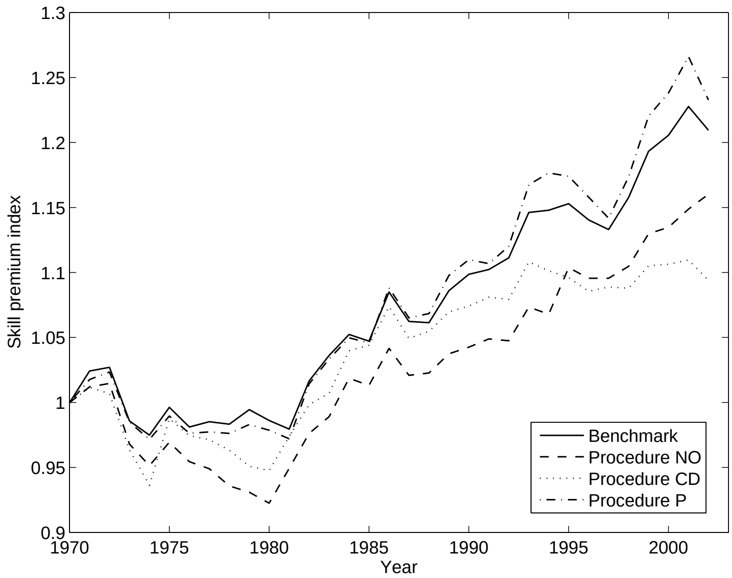

Figure 11: Skill premium in United States for various procedures to handle topcoded earners.

we assume that the missing observations 1971–1989 have the same time profile as the ratio of number of skilled to unskilled workers in the survey data. To remove large erratic swings in a few pairs of successive annual observations in the survey data, we use a Hodrick-Prescott filter with smoothing parameter 1. We obtain similar results by using a moving average filter with length five.

Appendix A.6 Variables in data sets

Table A1: CPS variables used for the U.S.

Variable label		Variable name
Person Weight		wgt
Age		age
Sex		sex
Race		_race
Educational attainment	< 1992	grdcom & _grdhi
	≥ 1992	grdatn
Employment status		esr (mlr for 1994)
Occupation Last Year,		occlyr
Industry Last Year,		indlyr
Class of Worker Last Year		_clslyr
Weeks Worked Last Year		_wkslyr
Hours worked	< 1976	hours
	≥ 1976	hrslyr
Wage and salary income	< 1989	_incwag
	≥ 1989	incer1, incwg1 & ernsrc
Allocated income		aincwag

Table A2: Census variables used for Sweden, 1970 and 1990

Variable label		Variable name
Age		Alder
Sex		Kon
Employment code	1970	AnstArtFoB70Ind
	1990	SektorFoB90
Educational attainment	1970	YrkeUtbAlmFoB70Ind & SkolUtbAlmFoB70Ind
	1990	SUN3FoB90
Industry	1970	SNI4FoB70
	1990	SNI2FoB90
Hours worked per week	1970	SyssFoB70Ind
	1990	ArbTidFoB90
Wage and salary income		ArbInk

Table A3: LOUISE variables used for Sweden, 1990–2002

Variable label		Variable name
Age		Alder
Sex		Kon
Employment code		YrkStalln
Educational attainment	< 2000	HUtbSUN
	≥ 2000	SUN2000Niva & SUN2000Inr
Industry	< 2002	AstSNI92
	2002	AstSni2002
Hours worked per week	1990	ArbTid
Wage and salary income		ArbInk

Appendix B KORV's model

Appendix B.1 Equilibrium properties

By solving for k_s from first-order condition (7) and substituting into production function (6), we get an equilibrium expression for output

$$y = (A\,\Omega)^{\frac{1}{1-\alpha}} \left[\frac{\alpha}{r_s}\right]^{\frac{\alpha}{1-\alpha}} \left\{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}}\right\}^{\frac{1}{\sigma}} . \tag{14}$$

From first-order conditions (8) and (10), we get an equilibrium expression for the quantity of equipment,

$$k_e = \left[\frac{w_s}{\theta_s r_e}\right]^{\frac{1}{1-\rho}} h_s \, . \tag{15}$$

From first-order conditions (9) and (10), we get an equilibrium expression for the labor share of output, χ,

$$\begin{aligned}\chi \equiv \frac{w_u h_u + w_s h_s}{A\,F(k_s,k_e,h_u,h_s)} &= (1-\alpha)\frac{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}}\,\theta_s h_s^\rho}{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}}} \\ &= (1-\alpha)\frac{1+\pi\,\dfrac{h_s}{h_u}}{1+\left[\pi^\sigma\,\theta_u^\rho\,\theta_s^{-\sigma}\left(\dfrac{h_s}{h_u}\right)^{\sigma(1-\rho)}\right]^{\frac{1}{\sigma-\rho}}},\end{aligned} \tag{16}$$

where $\pi \equiv w_s/w_u$ is the skill premium, and the last equality in (16) is obtained by twice invoking expression (13).

Our calibration procedure is motivated by the following observation on sets of equilibria and the choice of parameter values.

Claim 1. *For given labor inputs $\{h_s,h_u\}$ and rental rates $\{r_s,r_e\}$, suppose there exists an equilibrium $\{w_s, w_u = w_s/\pi, \chi, k_s, k_e\}$. Then by varying the initial parameters $\{\theta_u,\theta_s\}$, there exists a continuum of other equilibria, $\{\hat{w}_s, \hat{w}_u = \hat{w}_s/\pi, \chi, \hat{k}_s, \hat{k}_e\}$, where the skill premium, π, and the labor share of output, χ, are unchanged. In particular, for any skilled labor wage $\hat{w}_s \in (0,\infty)$, such an equilibrium is found by selecting the parameters $\hat{\theta}_u = (\hat{w}_s/w_s)^\sigma\,\theta_u$ and $\hat{\theta}_s = (\hat{w}_s/w_s)^\rho\,\theta_s$.*

This claim can be verified as follows. First, after invoking equilibrium expressions (14) and (15) for output and equipment, we confirm that first-order condition (10) for the employment of skilled labor continues to be satisfied at the given labor inputs $\{h_s,h_u\}$ under the alternative equilibrium wage $\hat{w}_s$ and parameter values $\{\hat{\theta}_u,\hat{\theta}_s\}$. Second, we can similarly confirm that first-order condition (9) for the employment of unskilled labor continues to hold under the alternative equilibrium wage $\hat{w}_u = \hat{w}_s/\pi$. Third, given the unchanged values of labor inputs and the skill premium, we confirm that the labor share of output in (16) is also unchanged, i.e., we verify that $\hat{\theta}_u^\rho\,\hat{\theta}_s^{-\sigma} = \theta_u^\rho\,\theta_s^{-\sigma}$.

Appendix B.2 Calibration of $\{\Omega, \theta_u, \theta_s\}$

For a given skilled labor share in 1970, rental rates, and parameters $\{\alpha, \sigma, \rho\}$, here is how we choose remaining parameters $\{\Omega, \theta_u, \theta_s\}$ to calibrate the model to a particular labor share of output (χ) and skill premium (π) in 1970, while allowing for any absolute wage level (w_s).

1. For an arbitrary value of Ω, we can compute the labor share of output as a function of the parameters θ_u and θ_s. Such a mapping is illustrated in Figure 12 for one particular value of Ω. Given a calibration value for the labor share of output, say $\chi = 2/3$, the bold curve in Figure 12 depicts pairs of parameter values (θ_u, θ_s) that produce the targeted value χ.

2. Both the labor share and the skill premium are constant along the bold curve in Figure 12, as described in Claim 1. Hence, given the calibration value of the labor share, the bold curve in Figure 12 represents a mapping from the parameter Ω (that is held fixed in the graph) to a specific value of the skill premium, π. Given the calibration value of the labor share, the compilation of pairs (Ω, π) from successive graphs when varying Ω, yields a mapping as illustrated in Figure 13.

3. Given a calibration value for the skill premium, Figure 13 pins down our choice of parameter Ω. Given that value of Ω, the associated Figure 12 contains a bold curve depicting permissible parameter values θ_u and θ_s. On the basis of Claim 1, we can find parameter values (θ_u, θ_s) to match any absolute wage level; we set $w_s = 1$ for both countries.

Note that the skill premium in 1970 would be identically equal to one if we had chosen 1970 group wages as weights in the construction of skilled and unskilled labor quantities. But since we have used average group wages for 1970–1990, each country's skill premium in 1970 differs slightly from being equal to one.

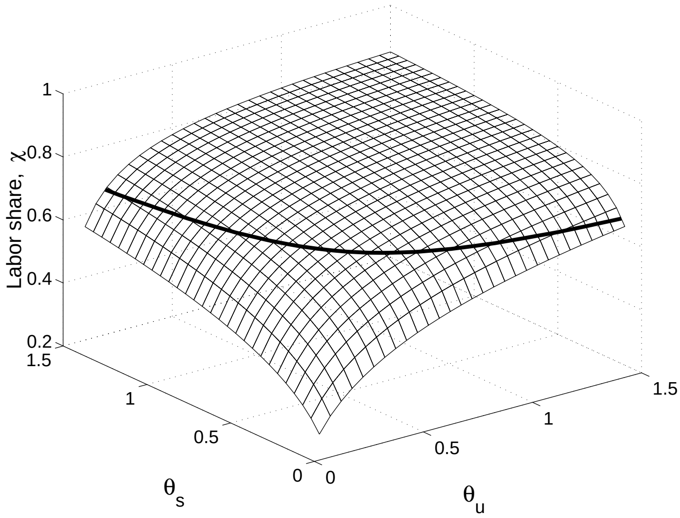

Figure 12: Labor share of output, χ, as a function of parameters θ_u and θ_s, for a given value of the parameter Ω. Along the bold curve is not only the labor share constant, $\chi = 2/3$, but so is the skill premium.

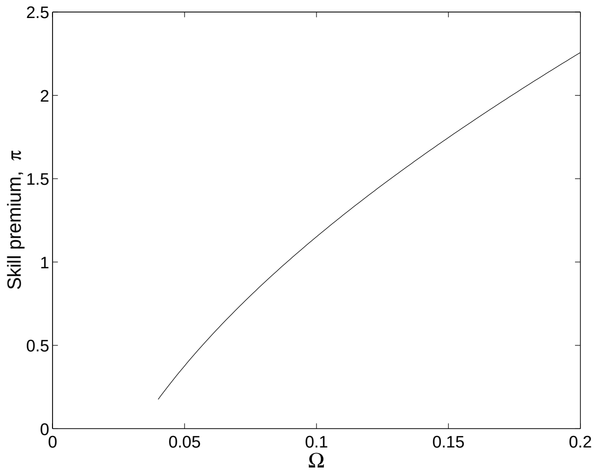

Figure 13: Skill premium, π, as a function of the parameter Ω, given that the labor share of output is held constant, $\chi = 2/3$.