The Missing Swedish Skill Premium:
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 around 9 percent. Since then both skill premia have increased by about 10 percentage points in 2002. A theory that equalizes 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 implement the welfare state.

Key words: Skill premium, employment, private sector, public sector, capital-skill complementarity, Sweden, United States.

JEL-classification: E24, J31.

1   Introduction

When it comes to trans-Atlantic differences in wage outcomes such as relative earnings of skilled versus unskilled labor, one can discern two persuasions among economists. One persuasion trusts that observed wages can be explained by workers' marginal products, while the other persuasion is skeptical and seeks alternative explanations. We quantitatively examine if a theory that equalizes wages with marginal products can explain outcomes in what is possibly the most challenging European country, Sweden, with its well-known small wage differentials and, hence, a low skill premium. First, we compute as accurately as possible a measure of the Swedish skill premium based on population data and second, we use a structural model to see if the skill premia in Sweden and the U.S. can be rationalized in terms of observed labor quantities between 1970-2002.

Using census data and tax records that cover the entire working-age population, we document a dramatic collapse of the Swedish skill premium that is more spectacular than earlier accounts have suggested. The skill premium fell by more than 30 percent between 1970 and 1990, and recovered by no more than 10 percentage points in 2002. In comparison, the U.S. skill premium, after an initial decline in the 1970s, rose by around 9 percent over the same period 1970-1990 and an additional 11 percentage points in 2002 based on the *Current Population Survey* (CPS). Following earlier studies, our computations divide 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.).

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.[1] This has led many observers to question if Swedish workers are any longer paid their marginal products. For example, Edin and Topel theorize that Swedish wages have become unhinged from marginal products because of collusion between the trade union and the employers' association.[2] In contrast, we explore if a theory that equalizes wages with marginal products can rationalize these disparate outcomes.

Our analytical framework is the production function with a capital-skill complementarity formulated by Krusell, Ohanian, Ríos-Rull and Violante (2000), hereafter denoted KORV, who showed that the U.S. skill premium between 1963 and 1992 can be accounted for by observed quantities of skilled and unskilled labor.[3] The historical increase in the relative supply of skilled labor is more than offset by skill-biased technological change, as measured

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[1] For a contemporary study that lends empirical support to such a high Swedish skill premium in the late 1960's, see Ståhl (1974) who computes rates of return to university education.

[2] In Edin and Topel's (1997, pp. 192–197) model, the egalitarian objectives of the trade union coalesce with the interests of the employers' association because 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."

[3] For another study that explains the U.S. skill premium in terms of observed aggregate labor quantities, see Katz and Murphy (1992). In section 2, we briefly refer to their study and its implications for Sweden.

by falling quality-adjusted equipment prices. We confirm that the model's explanatory power extends to 2002 for the U.S. economy, but we also show that total employment of skilled and unskilled labor in Sweden cannot account for the Swedish skill premium. Next, we ask if measures of private sector employment in Sweden can rationalize the dramatic evolution of the Swedish skill premium.

How could the division of total employment in private and public employment affect the skill premium? If the public sector produces perfect substitutes to the output of the private sector, using the same relative inputs of skilled and unskilled labor, there would be no effects of public sector employment on the skill premium. But public sector output often differs from what would be demanded and produced, if the workers were instead employed in the private sector; and even when its output is a close substitute to that of the private sector, the public sector will most likely neither attain the efficiency nor reflect the scale and composition of an alternative market-driven allocation in the private sector.[4] Hence, a distinction between public and private sector employment is warranted. In the simplest theory, the two sectors are assumed to produce different outputs, and free mobility of labor implies that workers of each type are paid the same across sectors. When private firms are competitive and maximize profits, labor is paid its marginal product in private employment. Under the assumption of constant returns to scale, it follows that the public sector can only affect the skill premium if it disproportionately hires one type of labor and thereby changes the relative composition of labor available to the private sector. For example, if the public sector disproportionately expands its employment of unskilled workers, it would instill greater scarcity of unskilled labor in the private sector which would cause the marginal product of unskilled labor to rise relative to that of skilled labor in private employment. Hence, a disproportionate expansion in public sector employment of unskilled labor would drive down the economy's skill premium.

We show that the relative supplies of skilled and unskilled labor in the private sector can indeed explain the evolution of the Swedish skill premium between 1970-2002.[5] The model attributes the dramatic decline of the Swedish skill premium to an expanding public sector that today comprises roughly one third of the labor force, and to the fact that this expansion has largely taken the form of drawing low-skilled workers into local government jobs that implement the welfare state. If the public sector in Sweden had not disproportionately hired unskilled workers, we show in a counterfactual experiment that the Swedish skill premium would not have fallen but rather increased in 2002 to a level similar to that of the U.S. Our analysis suggests that the causes to compressed skill premia in other European countries might be sought in the withdrawal of low-skilled workers from the private sector into public sector employment, long-term unemployment, or disability and early retirement programs.

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[4]An example of differences in public and private sector allocations is offered by Rosen (1996), who studied the public provision of child-care services in Sweden. He concluded that there is an excessive supply of such services in Sweden which "reduces the value of social output and living standards in the economy. Total output is smaller than it would have been if household services had been paid privately and transacted through the market"(Rosen, 1996, p. 740).

[5]The predictive power of the model for the U.S. is largely unaffected by the choice of either private or total employment.

The rest of the paper is organized as follows. Section 2 describes earlier studies that have sought to quantitatively explain the Swedish skill premium, but that were mired in data problems. Section 3 presents our data, and section 4 lays out KORV's model. The skill premia in the U.S. and Sweden are simulated in section 5 and additional implications of the model are examined. Section 6 interprets the Swedish experience through the lens of the model, while section 7 concludes with a discussion of our findings. Two appendices elaborate on the data and the theoretical framework.

2 Earlier studies of the Swedish skill premium

An early attempt to explain the Swedish skill premium in terms of aggregate labor inputs was made by Edin and Holmlund (1995) for the period 1970-1990. Inspired by Katz and Murphy (1992), they estimated the log of the skill premium as a function of the log of the ratio of skilled to unskilled labor, and a linear time trend that represents a relative demand shifter for skilled labor. In contrast to Katz and Murphy's (1992) findings for the U.S., Edin and Holmlund (1995, table 9.7) estimated a much smaller time trend (0.8 log points as compared to 3.3) and a much higher elasticity between skilled and unskilled labor (2.9 as compared to 1.4). As we now describe, these large differences can be attributed to the use of a poor proxy for the Swedish skill premium.

Lacking the type of data that enabled Katz and Murphy (1992) to compute quantity and wage indices for the two aggregate inputs of skilled and unskilled labor in the U.S., Edin and Holmlund (1995) relied on proxies in their Swedish study. Regarding the skill premium, they chose to "take the private-sector university [versus high school] wage premium as a proxy" and, specifically, they considered three alternative measures as shown in figure 1. The upper solid line in figure 1 depicts Edin and Holmlund's proxy of the skill premium that was used in their main estimation, which is the university wage premium according to Statistics Sweden's (SCB) monthly salary data for male white-collar workers in mining, manufacturing and construction.[6] The other two measures are deduced from year-specific Mincer equations based on household surveys, where proxy 2 (proxy 3) in figure 1 depicts the estimated wage differential between workers with 15 versus 12 (16 versus 12) years of education. Our question is if any one of these time series of the university wage premium represents a good proxy for the *aggregate* Swedish skill premium?

As described below, using census data and tax records for the entire Swedish labor force, we compute quantity and wage indices for the aggregate inputs of skilled and unskilled labor in Sweden, precisely as Katz and Murphy (1992) and KORV did for the U.S. The lower solid line in figure 1 depicts the resulting skill premium in Sweden. In hindsight, we see that the university wage premium according to proxy 3 in figure 1 is the proxy that is closest to capture the actual fall in the skill premium between 1970 and 1990. Interestingly enough, when Domeij and Ljungqvist (2006) estimate the Katz-Murphy equation for Sweden using the time profile of this proxy but where its total change between 1970 and 1990 is rescaled to

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[6]We thank Matthew Lindquist for providing us with the time series for proxy 1 in figure 1.
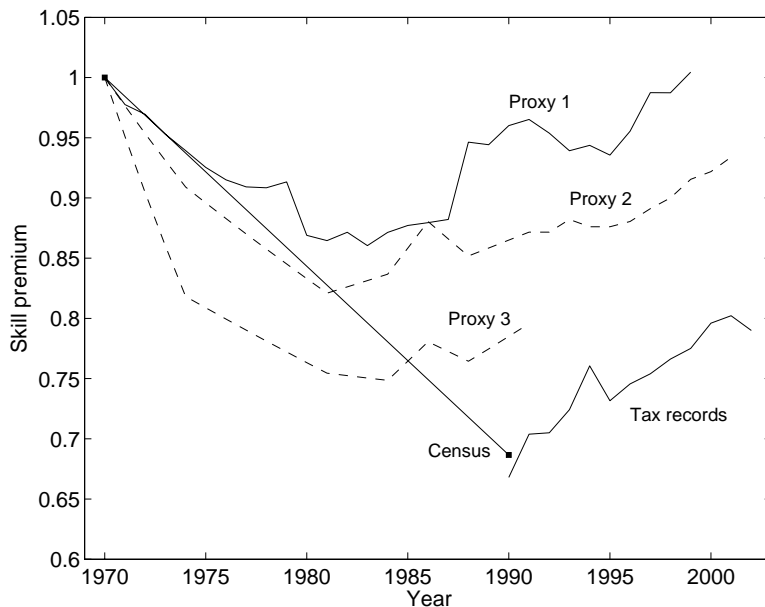
Figure 1: Proxies for the Swedish skill premium in private employment which were used in earlier studies (normalized to one in 1970). Proxy 1 is the university wage premium according to Statistics Sweden's (SCB) monthly salary data for male white-collar workers in mining, manufacturing and construction. Proxy 2 (proxy 3) was constructed by Edin and Holmlund (1995) who used household surveys and year-specific Mincer equations to estimate the wage differential between workers with 15 versus 12 (16 versus 12) years of education. Proxy 2 was extended to year 2001 by Gustavsson (2006). The lower solid line is the actual skill premium based on census data and tax records for the entire Swedish labor force, as described in section 3.2.

mimic the full 31% fall in the actual aggregate skill premium, the estimated coefficients are almost the same as those that Katz and Murphy (1992) obtained for the U.S. – the Swedish time trend is 3.0 log points and the elasticity of substitution between skilled and unskilled labor is 1.3 as compared to the U.S. numbers of 3.3 and 1.4, respectively.[7]

Encouraged by the Katz-Murphy regression results in Domeij and Ljungqvist (2006), the present paper examines a structural model to see if the evolution of the Swedish skill premium can be explained in terms of marginal products of labor. The advantage of a structural model is that additional implications of the analysis such as the labor share of income, investment rates for structures and capital, and the economy's growth rate can be

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[7]Following Edin and Holmlund (1995), Domeij and Ljungqvist (2006) construct a proxy for the ratio of skilled to unskilled labor between 1970 and 1990 by counting workers with and without a college education according to labor-market surveys (*Arbetskraftsundersökning*, AKU). We thank Per-Anders Edin for providing us with the data. As compared to the population measures in the present paper, the proxy does a fairly good job in capturing the relative change in the ratio of skilled to unskilled labor between 1970 and 1990.

compared to data. KORV have demonstrated that their structural model can explain the evolution of the U.S. skill premium, so we adopt the same framework and let their theory of a capital-skill complementarity enter as an auxiliary hypothesis in our study of the Swedish skill premium.

Another recent study that uses KORV's model to study the Swedish skill premium for the period 1970-1999 is Lindquist (2005). But once again, that study adopts the earlier poor proxy for the Swedish skill premium, namely, proxy 1 in figure 1. For example, while proxy 1 suggests that the aggregate Swedish skill premium in 1999 is back to its 1970 value, the actual skill premium in 1999, based on census data and tax records, is not anywhere near its historical level of inequality in 1970 but, rather, more than 20% depressed relative to 1970.

3   Data

As reviewed, earlier studies of the Swedish skill premium relied on proxies for the *aggregate* skill premium whereas a true measurement would involve the construction of wage indices for the composite of workers that make up the two aggregates of skilled and unskilled labor. This section presents our data and how the measures for skill premia and labor inputs are constructed. We use the aggregation methods of KORV, and Katz and Murphy (1992). For a more complete description, see Appendix A.

3.1   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.

3.2   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.[8] Hence, for the years 1970 and 1990-2002, we obtain almost perfect measures of the skill premium and of the relative supply of skilled and unskilled labor.[9]

3.3   Trends in skill premia and labor supply

We lay out the basic facts in figures 2 through 4. Figure 2 shows how the Swedish and U.S. skill premia evolved between 1970 and 2002.[10] The U.S. development is well known. The skill premium deteriorated during the 1970s, but has grown rapidly since around 1980.[11] The development of the Swedish skill premia is dramatically different. Between 1970 and 1990 the skill premium fell by an astonishing 31 percent. Thereafter it recovered somewhat, but in 2002 it was still 20 percent lower than in 1970.

   Figure 3 shows strong growth in the relative supply of college educated workers in both countries. Between 1970 and 2002 the relative supply of skilled workers rose by 146 and 240 percent in the U.S. and Sweden, respectively. The difference across countries is further accentuated if we separate out private employment.[12] Figure 4 depicts the ratio of skilled/unskilled labor in private employment relative to that in total employment. While this quotient has gone up by around 20 percent in the U.S. between 1970 and 2002, the Swedish increase has been almost threefold of that. Consequently, the relative supply of

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

[9]The only shortcoming of the Swedish data is that while it contains individual data on hours worked for 1970 and 1990, we have to extrapolate that information for the years 1991-2002. Specifically, we assume that average hours of work in each group of workers has remained constant at its 1990 level. This is consistent with Statistics Sweden's estimates of almost constant hours worked by age and gender among the employed between 1990 and 2000 and hence, our extrapolations are not likely to undermine the advantage of otherwise having full information about the entire population of Sweden. See Appendix A.3 for more details.

[10]Regarding all Swedish time series, data from the Census of Population Surveys for 1970 and 1990 are marked by bullets and connected by straight lines. The minor difference between a bullet in 1990 and the corresponding curve starting in 1990 and going forward reflects a slight change in the classification of education in the LOUISE database as compared to the census data.

[11]Our time series for the skill premium in the U.S. is very similar to that documented by Autor et al. (2008). 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 data is topcoded.

[12]We define the difference between total and private employment as being those employed in health, education, postal services and government administration. This delineation is appropriate for Sweden where these services are almost exclusively produced in the public sector. Regarding the U.S., we keep the same definition for comparative purposes while emphasizing that the quantitative analysis of the U.S. skill premium is largely unaffected whether we use total employment or our definition of private employment. For the U.S. we have also used the Census of Population Survey's classification into private and public sector employment with very similar results.
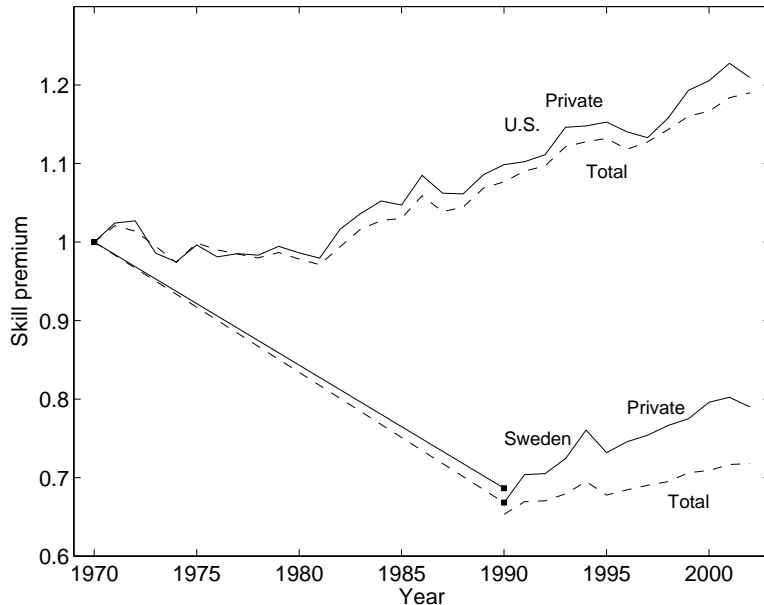
Figure 2: Skill premium in Sweden and the U.S. (normalized to one in 1970). The solid lines refer to private employment, and the dashed lines to total employment.

college educated workers in Swedish private employment has grown much more than that in total employment in figure 3. As we describe in section 6.1, the driving force has been a rapid expansion of the Swedish public sector that has disproportionately hired unskilled labor.

Our task in this paper is to take the labor supply series in figures 3 and 4 as given, and then explain the skill premium series in figure 2.

4   Model

We analyze the skill premium in terms of the production function proposed by KORV,

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

where y is aggregate output, k_s and k_e are the stocks of capital structures and capital equipment, and h_u and h_s are hours of unskilled and skilled labor inputs. The only potentially time varying production parameter is A, i.e., a multiplicative neutral technology factor. The remaining time invariant production parameters are as follows. Capital structures share of income is given by α, and σ and ρ determine elasticities of substitution. The elasticity of substitution between skilled labor and capital equipment is $1/(1 - \rho)$, and the elasticity
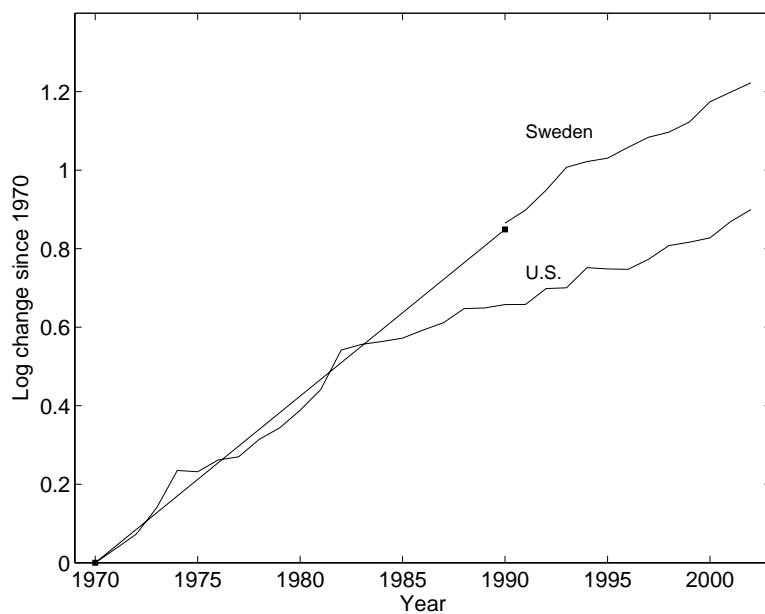
Figure 3: Ratio of skilled to unskilled labor in total employment in Sweden and the U.S. (log change since 1970).
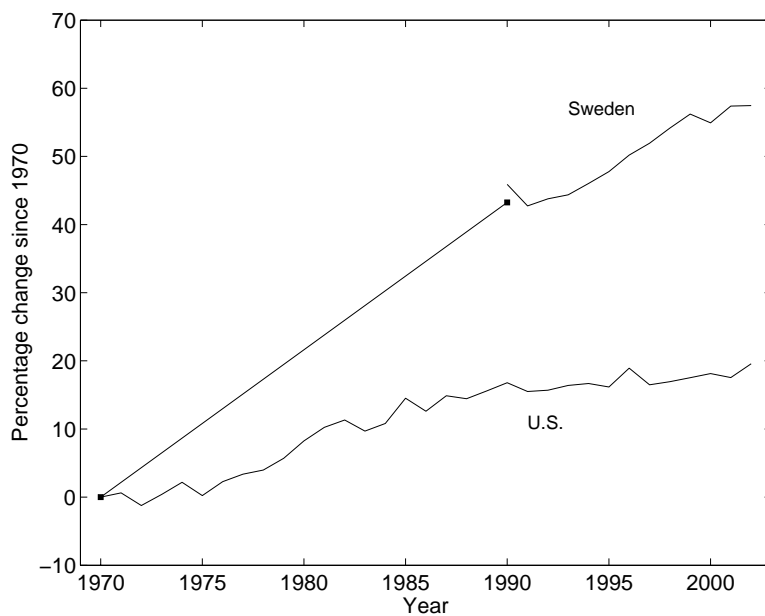

Figure 4: Ratio of skilled/unskilled labor in private employment relative to that ratio in total employment in Sweden and the U.S. (percentage change since 1970).

between unskilled labor and equipment or skilled labor is $1/(1-\sigma)$.[13] Income shares of different production factors are governed by μ and λ, and ψ_u and ψ_s convert hours of unskilled and skilled labor, respectively, into efficiency units.

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 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.

KORV note that one of the parameters $\{\mu, \lambda, \psi_u, \psi_s\}$ must be fixed as a normalization in their estimation. To make this point explicit, we map the four parameters into three, $\{\Omega, \theta_u, \theta_s\}$, by rewriting production function (1) as follows,

$$y = A\,\Omega\,k_s^\alpha \left\{ \theta_u\,h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}} \right\}^{\frac{1-\alpha}{\sigma}} \equiv A\,F(k_s, k_e, h_u, h_s), \tag{2}$$

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

$$r_s = A\,\alpha\,F(k_s, k_e, h_u, h_s)\,k_s^{-1}, \tag{3}$$

$$r_e = A\,\Gamma(k_s, k_e, h_u, h_s)\,(k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}}\,k_e^{\rho-1}, \tag{4}$$

$$w_u = A\,\Gamma(k_s, k_e, h_u, h_s)\,\theta_u h_u^{\sigma-1}, \tag{5}$$

$$w_s = A\,\Gamma(k_s, k_e, h_u, h_s)\,(k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}}\,\theta_s h_s^{\rho-1}, \tag{6}$$

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

$$r_s = r + \delta_s, \tag{7}$$

$$\frac{r_e}{p_e} = 1 + r - \frac{p'_e}{p_e}(1-\delta_e) \equiv \tilde{r}_e, \tag{8}$$

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[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.

[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.

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

From first-order conditions (5) and (6), we get an equilibrium expression for the skill premium,

$$\frac{w_s}{w_u} = \left(k_e^\rho + \theta_s h_s^\rho\right)^{\frac{\sigma-\rho}{\rho}} \frac{\theta_s}{\theta_u} \frac{h_s^{\rho-1}}{h_u^{\sigma-1}}. \tag{9}$$

Appendix B.1 further characterizes equilibrium properties of KORV's model.

4.1 Calibration procedure

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 (3) and (4). (The implied investment rates provide an additional dimension for assessing the performance 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 (9).

The simulations are carried out in models that are calibrated to a base year (1970 in the benchmark calibrations). Besides outcomes in that base year, our calibration procedure draws upon observations in the literature on the average annual real return (r), depreciation rates (δ_s, δ_e), and equipment price movements (p'_e/p_e), which are assumed to be the same across countries. We use country-specific historical averages of the labor income share and the capital structures share of income (α). To complete the parameterization of the U.S. model, we use KORV's estimates for the elasticity parameters σ and ρ. In the absence of comparative Swedish estimates, we need to deduce those from our data and model, as explained in section 4.3.

For further details about the calibration procedure, see Appendix B.2.[15]

4.2 Parameterization common to Sweden and the U.S.

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 Sweden and the U.S. share the following identical premises.

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

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[15]Appendix B.2 describes how the remaining scale parameters, $\{\Omega, \theta_u, \theta_s\}$, are calibrated based on outcomes in the base year.

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.

    b) 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. 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.

    c) 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 a)–c) yield rental rates on structures, $r_s = 0.09$, and equipment, $\tilde{r}_e = 0.209$, as given by equations (7) and (8). Since premises a)–c) are assumed to be the same across Sweden and the U.S., it follows that these rental rates are also the same.

4.3   Country-specific parameter values

To complete the parameterization of the U.S. model, we set the labor share of income equal to 2/3, which is a common value in quantitative studies of the U.S. economy. We use KORV's estimates of parameters $\alpha = 0.117$, $\sigma = 0.401$ and $\rho = -0.495$, i.e., capital structures' share of income is 11.7 percent, and the substitution elasticities are 1.67 between unskilled and skilled labor $(1/(1-\sigma))$, and 0.67 between skilled labor and equipment $(1/(1-\rho))$.

    In contrast to the common parameterization of section 4.2, we allow income shares and substitution elasticities to differ across countries. An ideal theory might have ensured that all primitives in form production functions and preferences could be calibrated once and for all, and then be expected to hold across different economies and time periods. But this is usually not the case in quantitative macroeconomics and our study is not an exception.[16]

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[16]As econometricians reestimate regression equations based on different data sets, applied macroeconomists routinely recalibrate their models. For example, in *Great Depressions in the Twentieth Century* edited by Kehoe and Prescott (2007), different calibrations of the same real business cycle model are employed for the U.S., Canada, France, Germany, Argentina, Mexico, Chile, Japan, Brazil, New Zeeland, Switzerland, and Finland, respectively.

To complete the parameterization of the Swedish model, we use historical averages to estimate the capital structures' share of income, $\alpha = 0.17$ in both the private sector and in the total economy; and a labor share of income equal to 0.61 in the private sector and 0.68 in the total economy. Sensitivity analyses inform us that our conclusions are quite robust to the choice of these factor income shares.[17] Regarding elasticity parameters σ and ρ, we lack good estimates for Sweden and we choose instead to deduce these from our data and model. In particular, we split the Swedish data in two subperiods; 1970–1990 and 1990–2002. We use data from one subperiod to calibrate the elasticities of the model, and data from the other subperiod to assess the model's predictive power.[18] That is, we perform two calibration exercises for Sweden with associated 'out-of-sample' predictions. To anticipate our results, we find that private employment performs well in both exercises while total employment fails to predict the evolution of the Swedish skill premium.

5 Simulations 1970–2002

In this section, we first show that the explanatory power of KORV's earlier analysis of the U.S. skill premium extends to yet another decade up to 2002, and we provide a brief comparison between our calibration procedure and KORV's original estimation approach. We then apply KORV's model to Sweden, and we find that private employment, but not total employment, can explain the evolution of the Swedish skill premium. The section concludes with a look at other implications of the simulations, all of which lend empirical support to the theoretical framework.

5.1 U.S. skill premium

We use the calibrated model and the observed time series for the labor skill composition over the period 1970-2002 to predict the U.S. skill premium. Figure 5 shows that the predicted skill premium is remarkably close to the actual evolution, especially given that we have used a single year from our data set to calibrate the model: the base year 1970. We conclude that the explanatory power of KORV's earlier analysis of the U.S. extends to yet another decade up to 2002 and, it makes not much of a difference whether we use private or total employment. In contrast, we will show that private, but not total, employment can explain the evolution of the Swedish skill premium.

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[17]For details on the computations of factor income shares in Sweden, see Appendix A.5. As a sensitivity analysis we have set $\alpha = 0.1$ or $\alpha = 0.2$, and our results were largely unchanged. We have also set the labor income share to 2/3 as for the U.S. economy. Again, our findings were robust. These and the other sensitivity analyses in the paper are available from the authors upon request.

[18]Our finding of a somewhat lower elasticity of substitution between unskilled and skilled labor in Sweden as compared to that in the U.S. is consistent with the microeconomic study of Bergström and Panas (1992) who estimate relatively low elasticities in most industries of their Swedish data set. For our preferred calibration of the aggregate elasticity, see table 1.
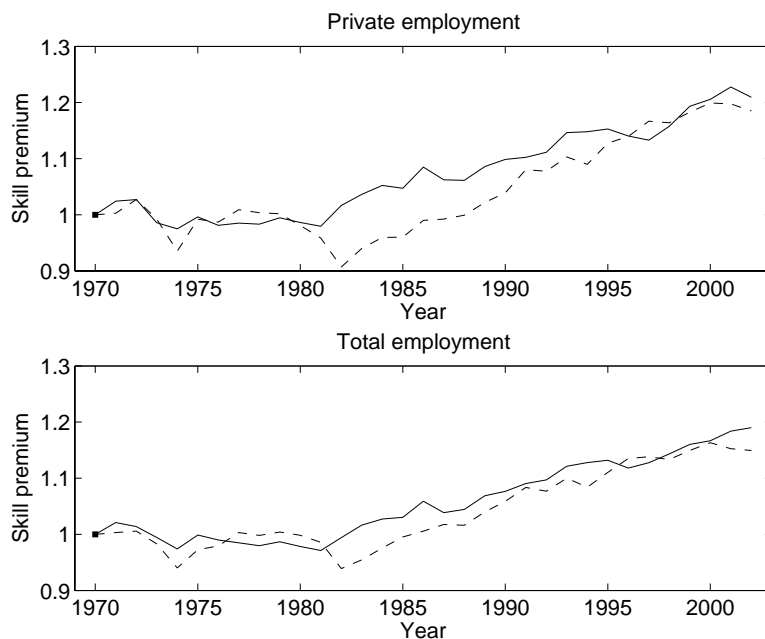
Figure 5: U.S. skill premium in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

5.1.1 A comparison with KORV's estimation approach

Since our simulation builds on the framework of KORV, it is useful to briefly describe how our analysis differs from theirs in three respects. 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. model to base year 1970 which ensures that our model perfectly explains the skill premium in that year. Second, except for observations on the labor skill composition over the period 1970-2002, our simulation does not draw on any other time series such as data on capital stocks that are used by KORV. Our simulation can be likened to a balanced-growth path perturbed by variations in the labor skill composition. The balanced-growth part of our simulation is driven by the premises of a constant rate of decline in equipment prices and a constant rate of return on investments. Third, in addition to the skill premium, our simulation predicts stocks of structures and equipment. In contrast to KORV who use capital stocks as inputs in their estimation, we will compare the implied investment rates from our simulation to those in the data to check the validity of our model.

5.2 Swedish skill premium

While our calibration of the U.S. model used a single year of data, we need more data points to calibrate the Swedish model. Specifically, we need to deduce estimates for the elastic-

ity parameters. The decline in the Swedish skill premium between 1970 and 1990, and its subsequent partial recovery over the next decade, suggest a natural calibration-simulation procedure. We use the first subperiod 1970–1990 (with base year 1970) to calibrate the elasticities, and then simulate the calibrated model over the entire period 1970–2002, including the 'out-of-sample' subperiod 1990–2002. Alternatively, as a robustness check, we reverse the order of the calibration-simulation procedure by using the second subperiod 1990–2002 (with base year 2002) to calibrate the elasticities, and then simulate the model over the entire period 1970–2002, including the out-of-sample subperiod 1970–1990. Besides these two out-of-sample simulation procedures, we use all data (with base year 1970) to calibrate the elasticities and then simulate the skill premium. If the theory is correct, the model should accurately predict the skill premium in all three procedures.

For each calibration-simulation procedure and for each measure of employment (private or total employment), we present a sensitivity analysis by considering different pairs (σ, ρ). In particular, for each ρ that belongs to a set that we have picked, we compute the value of σ that best captures the observed skill premium in the calibration period (using the least-squares metric). Given each such pair (σ, ρ), we then simulate the model over the whole period 1970–2002.

We pick the set of ρ so that the corresponding elasticities of substitution between skilled labor and equipment spans the interval $[0.3, 0.8]$. As we will show, this interval is enough to give both private and total employment the best chances of predicting the observed skill premium. Note also that the interval falls well within the range of empirical estimates, as reviewed by Hamermesh (1993).

5.2.1 Private versus total employment

The upper panel of figure 6 shows a remarkable good fit between simulated and actual skill premia when using private employment in Sweden. The calibration period is 1970–1990, i.e., substitution elasticities are chosen to capture the dramatic fall in the skill premium between 1970 and 1990. Hence, the model's success in explaining the two observations on the skill premium in 1970 and 1990 is to be expected. But since no data after 1990 enters the calibration procedure, the good fit over the 'out-of-sample' subperiod 1990–2002 is truly astonishing. In the lower panel of figure 6, the model based on total employment fails the test when it counterfactually predicts that the skill premium continues to decline after the initial large fall in the calibration period.

The reverse calibration-simulation procedure in figure 7 conveys a similar message. It is remarkable that with substitution elasticities calibrated to the second subperiod 1990–2002, private employment in the upper panel can retrace the dramatic decline in the Swedish skill premium in the out-of-sample subperiod 1970–1990. In contrast, and once again, the model based on total employment in the lower panel utterly fails to predict the reversal in the evolution of the skill premium. The performance of the latter model based on total employment cannot be improved upon by considering elasticities outside of the stipulated
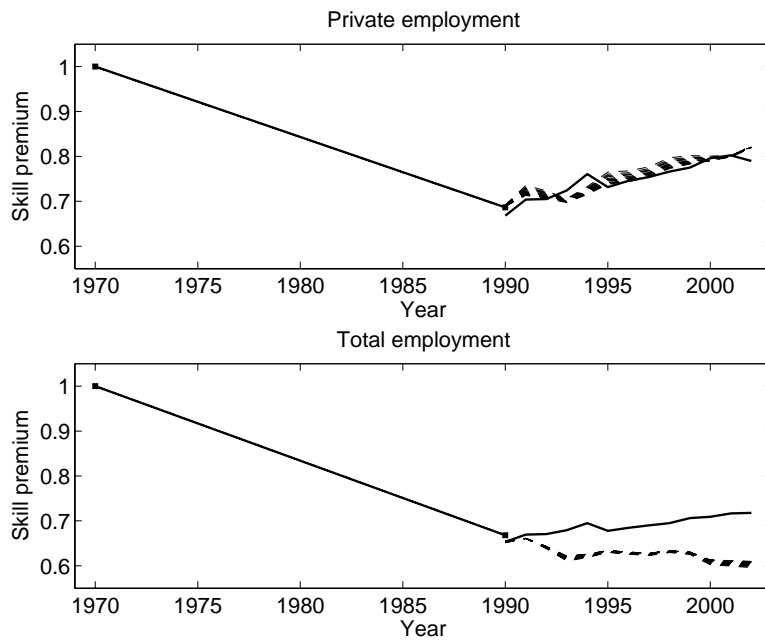
Figure 6: Swedish skill premium, given calibration period 1970-90, in private and total employment. The solid lines refer to data, and the dashed lines depict simulations.
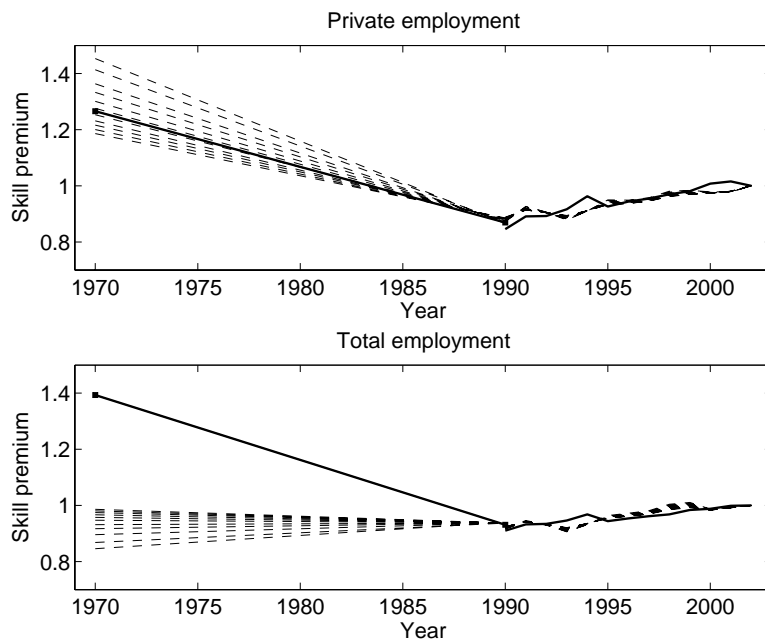

Figure 7: Swedish skill premium, given calibration period 1990-2002, in private and total employment. The solid lines refer to data, and the dashed lines depict simulations.
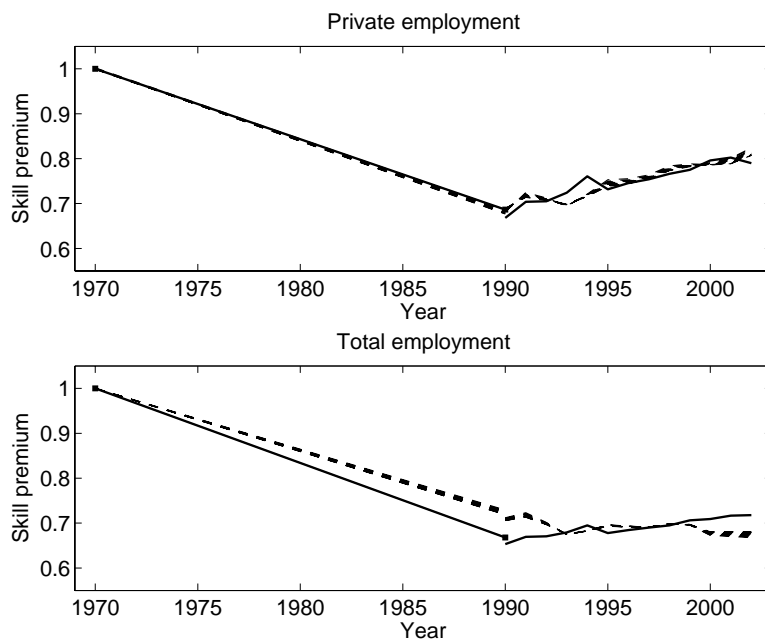
16

Figure 8: Swedish skill premium, given calibration period 1970-2002, in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

interval.[19]

In figure 8, we use the whole period 1970–2002 to calibrate the substitution elasticities. As compared to figure 6, the upper panel based on private employment in figure 8 does not look much different. That is, the additional data points in the calibration period cannot much improve upon what was already an almost perfect explanation of the skill premium. Concerning the lower panel based on total employment in figure 8, the path of the simulated skill premium can be understood in terms of the above out-of-sample calibration-simulation procedures. As shown in the lower panels of figures 6 and 7, total employment cannot explain the reversal in the skill premium evolution between the first and second subperiod. The enlargement of the calibration period to the whole period 1970–2002 does not change this fact. In particular, the lower panels of figures 6 and 8 are qualitatively the same – the

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[19]Since the upper simulated path in the lower panel of figure 7 predicts a skill premium in 1970 that is equal to its 2002 value, one might ask if the endpoint in 1970 could not be bent further upward by going outside of our stipulated interval of elasticities, $(1 - \rho)^{-1} \in [0.3, 0.8]$. Specifically, since a higher elasticity in the interval yields a higher skill premium in 1970, why not consider a yet higher elasticity than 0.8? This will not be helpful for two reasons. First, further increases in the elasticity hardly change the skill premium in 1970, as suggested by fact that the simulated paths are getting ever closer to each other at the upper echelon (where the paths correpond to evenly spaced elasticities). Second, further increases in the elasticity imply an income share of labor that strays ever farther away from data. In particular, the depicted upper path is associated with an income share of labor that strays 10 percent from its historical average.

skill premium is predicted to fall throughout the period 1970–2002. Unable to explain the reversal in the skill premium evolution, the enlargement of the calibration period in the latter figure can only improve upon the fit of the simulated skill premium by 'lifting up' the counterfactual persistent downward trend, and thereby reducing the least squares deviations from the true non-monotone path.[20]

5.3 Implications for labor share, productivity and investment

We have shown that private employment can explain the skill premium in both the U.S. and Sweden, but how do our models fare in other dimensions? Table 1 contains several key implications derived from our model simulation for the U.S., and for one of the Swedish calibrations in the upper panel of figure 8. Under the auxiliary assumption of no neutral technological growth, i.e., a time invariant parameter A in production function (1), we want to demonstrate that these other implications are also consistent with data.

Table 1: Calibration and simulations

	U.S.	Sweden
Country-specific calibration		
elasticity of substitution between skilled and unskilled labor	1.67	1.21
elastitity of substitution between skilled labor and equipment	0.67	0.35
capital structures share of income	0.117	0.170
labor share of income (in the base year 1970)	0.667	0.610
Simulated statistics 1970-2002		
average labor share of income	0.644	0.601
annual productivity gain[a]	2.0%	2.0%
annual growth rate of equipment (in efficiency units)	7.7%	7.3%

[a] Measured 'per hour,' i.e., we normalize $h_u + h_s = 1$ in each year.

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[20]The base year in figure 8 is 1970. Our results are unchanged if we use any other year as the base year, i.e, private employment can well explain the reversal in the skill premium evolution while total employment cannot. Consequently, the sum of squared deviations between simulated and actual skill premia is lower for private than for total employment, with the only exception being if we pick 1994 as the base year. Though, we would argue that private employment gives a more accurate picture even in that latter case for the following reason. Recall that, by construction, the simulated path perfectly explains the skill premium in the base year. Because there is an upward blip in the actual skill premium in 1994, it turns out that the correctly U-shaped simulated path based on private employment is lifted up above the actual path over the whole period, while the counterfactually persistent downward trend based on total employment happens to produce a slightly lower sum of squared deviations.

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The first implication in table 1 concerns the labor share of income. As noted by KORV, the specification of the production function does not guarantee a constant labor share. But the model for the U.S. yields a labor share that remains at roughly two-thirds of income throughout the period. For Sweden the model similarly predicts a stable labor share close to its historical average of 0.61.

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).[21] The model predicts that output per hour grew by 2 percent per year in both Sweden and the U.S., which is in line with what is observed in the data.[22]

The third implication concerns the growth rate of equipment. The model for the U.S. predicts that the stock of equipment in efficiency units grew at an annual rate of 7.7 percent which is very similar to the 7.4 percent reported by KORV based on Gordon's (1990) capital series. The model for Sweden suggests that Swedish equipment grew at the same pace. 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. National Income and Product Accounts. KORV reports that U.S. equipment, as measured in the National Income and Product Accounts, 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.[23] While not conclusive, this suggests that the stocks of equipment developed similarly in both countries.

6 Sweden through the lens of our model

The dramatic decline in the Swedish skill premium between 1970 and 1990, and the partial recovery in the 1990s correspond to economic events that can be interpreted in terms of our model. Besides looking at the Swedish experience through the lens of our model, we ask a counterfactual question; what would have happened if the Swedish public sector had not disproportionately hired unskilled labor?

6.1 Public sector expansion in the 1970s and 1980s

In the Census of Population Surveys, public sector employment increased by 640,710 people between 1970 and 1990 while private sector employment remained essentially constant. Hence, our data is consistent with Lindbeck's (1997, pp. 1280, 1311) account of the Swedish

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[21] 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.

[22] Using data from the Groningen Growth and Development Center (http://www.ggdc.net, the Total Economy Database, ted07I.xls) the average growth rate in real GDP per hour between 1970 and 2002 was 1.9 percent in Sweden and 1.6 percent in the U.S.

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

employment experience, when he 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" labor.

Through the lens of our model, a critical assertion by Lindbeck is that the public sector has particularly bolstered demand for labor at the lower skill spectrum. Indeed, the sharply increasing ratio of skilled to unskilled labor in private employment relative to that in total employment between 1970 and 1990 in figure 4, implies a sharply increasing ratio of unskilled to skilled labor in public employment. Using household survey data, Björklund and Freeman (1997, pp. 62-63) document an associated disproportionate increase from 1968 to 1991 in the fraction of workers in the lowest quartile in the hourly earnings distribution who are employed in the public sector. They emphasize that "[t]he pattern in most countries is for low-skilled workers to be underrepresented in the public sector, and finding equal representation [in 1991] is therefore surprising. ...We interpret the upward trend and their proportionate representation in Sweden as signs that the public sector has, indeed, been a greater demand of their labor than in other countries."

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 (1996, pp. 733, 731), "Sweden has 'monetized' the household sector of its economy. ...the welfare state encourages excessive production of household goods and discourages production of material goods. Too many people provide paid household (family) services for other people in the subsidized state sector and not enough are employed in the production of material goods. This is what explains the employment statistics and the growth of local government employment in Sweden." Our analysis shows how the forces identified by Rosen have also had equilibrium repercussions on the Swedish skill premium.

6.2   Crisis and public sector stagnation in the 1990s

Sweden went through a major economic crisis in the early 1990s. According to Lindbeck (1997, p. 1304), "[b]y far the most spectacular macroeconomic development in Sweden in the early 1990s ...was the emergence of the deepest recession since the 1930s. The accumulated fall in GNP was 4 percent during the three-year period 1991-1993". Because of the resulting deficit in public finances, public sector employment "was reduced by about ten percent in the early 1990s. ...it became clear that public sector employment cannot increase forever, as a share of total employment, without undesired consequences for the allocations of resources and economic incentives" (Lindbeck, 1997, p. 1311).

Through the lens of our model, none of these events would have any effects on the Swedish skill premium unless they affected the evolution of the ratio of skilled to unskilled labor in private sector employment. After computing average annual growth rates of the

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ratio of skilled to unskilled labor in total employment and private employment, respectively, we conclude that there were indeed changes in this dimension between the periods 1970-1990 and 1990-2002. Specifically, while the annual growth rate of the ratio of skilled to unskilled labor in total employment slowed from 4.3% to 3.0%, the corresponding change in private employment was larger from 6.2% to 3.7%. Thus, the stagnating public sector in the 1990s did not absorb unskilled labor to the same extent as in the previous two decades. Observers of the Swedish labor market who have focused on measures of total employment have not registered this detrimental development for unskilled labor. The theory informs us that unskilled workers are better off when the public sector increases its demand for their services, because the ensuing relative scarcity of unskilled workers in the private sector drives up their marginal product and wage. Our analysis attributes the reversal in the 1990s from a falling to an increasing Swedish skill premium to two factors: (1) the Swedish growth in the ratio of skilled to unskilled labor in total employment slowed after a couple of years into the 1990s, as shown in figure 3; and (2) the disproportionate hiring of unskilled labor in the Swedish public sector weakened in the 1990s, as reflected in figure 4. Next, we will quantify how much the Swedish skill premium was affected by the fact that the public sector has favored unskilled labor.

6.3 What if the public sector had not favored unskilled labor?

We use our calibrated Swedish model in table 1 to answer the following counterfactual question. Suppose that the relative growth rate of the ratio of skilled/unskilled labor in Swedish private employment as compared to this ratio for total employment had been the same as that of the U.S. over the period 1970-2002, what would have been the Swedish skill premium in such a scenario where the public sector had not favored unskilled labor? Specifically, we keep the actual time series for the ratio of skilled to unskilled labor in Swedish total employment, but let post-1970 changes in the ratio of skilled/unskilled labor in private employment relative to that of total employment be governed by the corresponding evolution in the U.S.[24]

Figure 9 shows that the Swedish skill premium would have remained essentially flat between 1970 and 1990 if it had not been for the expansion of public sector employment which disproportionately hired unskilled labor. Moreover, moving into the 1990s with a slowing growth in the ratio of skilled to unskilled labor in Swedish total employment (the upper line in figure 3) coupled with an unchanging relative allocation of skilled and unskilled labor in private versus total employment (when assuming the U.S. evolution as given by the

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[24]The counterfactual ratio of skilled to unskilled labor in Swedish private employment in year t is computed as

$$\left(\frac{h_{st}^{\text{priv Swed}}}{h_{ut}^{\text{priv Swed}}}\right)^{\text{Counterfactual}} = \frac{\kappa_t^{\text{US}}}{\kappa_{1970}^{\text{US}}} \; \kappa_{1970}^{\text{Swed}} \; \frac{h_{st}^{\text{tot Swed}}}{h_{ut}^{\text{tot Swed}}} \qquad \text{where} \qquad \kappa_t^{\text{j}} \equiv \frac{h_{st}^{\text{priv j}}}{h_{ut}^{\text{priv j}}} \bigg/ \frac{h_{st}^{\text{tot j}}}{h_{ut}^{\text{tot j}}} \, ,$$

and h_{it}^{xj} is the labor quantity of skill type i (i = u[nskilled] or i = s[killed]) in employment x (x = priv[ate employment] or x = tot[al employment]) in country j (j = Swed[en] or j = US).
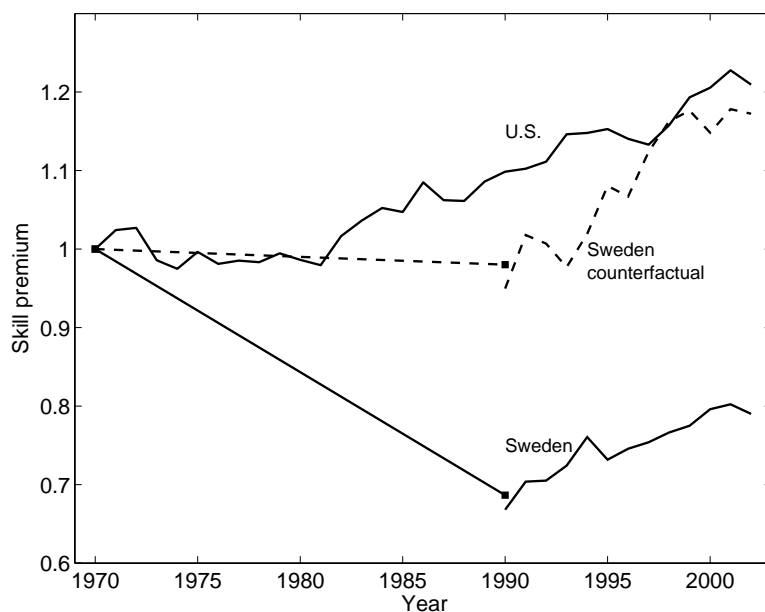
Figure 9: Skill premium in private employment in Sweden and the U.S. (normalized to one in 1970). The solid lines depict the data, while the dashed line for Sweden refers to a counterfactual experiment in which the evolution of the Swedish ratio of skilled/unskilled labor in private employment relative to that of total employment is assumed to be governed by the relative changes observed in the U.S.

lower line in figure 4), our counterfactual scenario results in a Swedish skill premium in 2002 that is not much different from that of the U.S. Given this rather remarkable prediction, we conclude that public sector employment in Sweden has been critical for suppressing the Swedish skill premium over the period 1970-2002 as compared to the U.S. skill premium.

7   Discussion

We are not the first ones to suggest that public sector employment can be an important determinant for wage outcomes.[25] 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.

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[25]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.)

7.1 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 partly a theory about nonmarket outcomes. On the one hand, 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. On the other hand, and in contrast to some other nonmarket theories, unskilled and skilled labor in our analysis are paid their marginal products in the private sector.

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 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 employed at wages equal to the marginal products of unskilled and skilled labor.

Our theory cannot discriminate between these alternative scenarios because, like KORV, we take the quantities of unskilled and skilled labor as given. Moreover, like that earlier study, 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.[26] 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 between 1970 and 1990, as we have documented with census data and tax records for the economy's entire population.

7.2 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

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[26]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.

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 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 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 globalization 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 the generous welfare programs of Europe 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.[27] A hypothesis that merits further investigation.

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[27] Our hypothesis is related to the 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). Blanchard (2004) reviews evidence that suggests as much as 6 percent of average French labor productivity is due to low-productivity workers having left the labor force. Those calculations are based on the extrapolation of a truncated wage distribution but do not take into account that the wage structure itself would be affected if skilled and unskilled labor are imperfect substitutes.

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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 Domeij and Ljungqvist (2006).

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 51,514 in 1970, 74,851 in 1980, 67,732 in 1990 and 94,371 in 2002.

In the wage sample we follow Autor et al. (2008) 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.[28] 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. We have also used the Census of Population Survey's classification into private and public sector employment with very similar results.

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.[29] Like Autor et al. (2008), 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

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[28]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.

[29]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. (2008) and reassign primary and secondary earnings to their topcoded values.

regard both types of graduates as high school graduates. 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 v_{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

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

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

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

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

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 $h_{jt} = \sum_{g \in G_j} H_{gt} \bar{W}_g v_{gt}$, where the weights $\bar{W}_g$ are the average group wages for 1970 and 1990;[30] and the class wage rate is calculated as $w_{jt} = \sum_{g \in G_j} H_{gt} W_{gt} v_{gt} \ / \ h_{jt}$.

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 ($v_{it} = 1$). 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

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[30] 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

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

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 $\geq$35 hours per week, and for 1990 the intervals are 1-15, 16-19, 20-34, $\geq$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.[31] 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 employed men remained constant while it increased by 3 percent for employed women. In the age dimension, only a small fraction of the labor force experienced large changes; individuals age 19 and younger. To investigate the importance of the assumption of constant hours in each group g, we perform the following robustness check. Average hours in each group g is adjusted by a time specific age$\times$gender factor such that average hours worked by age$\times$gender groups are consistent with Statistics Sweden's estimate for each year in the period 1991 to 2002. The results of the paper were not affected.

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[31] 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.
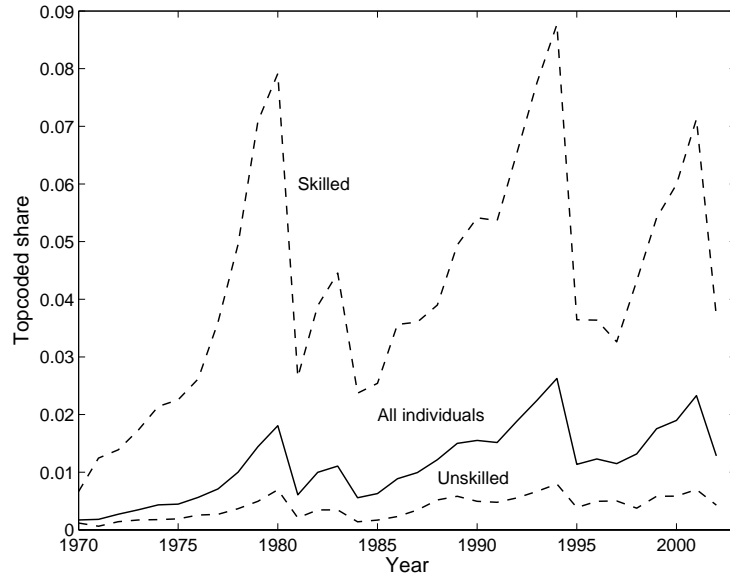
Figure 10: Share of individuals who are topcoded. The solid line depicts the topcoded share among all individuals while the upper (lower) dashed line shows the topcoded share among skilled (unskilled) labor. The topcode classification is based on total earnings until 1988, and primary earnings afterwards.

Appendix A.4   Topcoding in U.S. data

As mentioned above, data on wage and salary income is topcoded in U.S. data. Figure 10 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 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,[32] 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

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[32]For a recent study using the Pareto distribution to interpolate income distributions, see Piketty and Saez (2003).
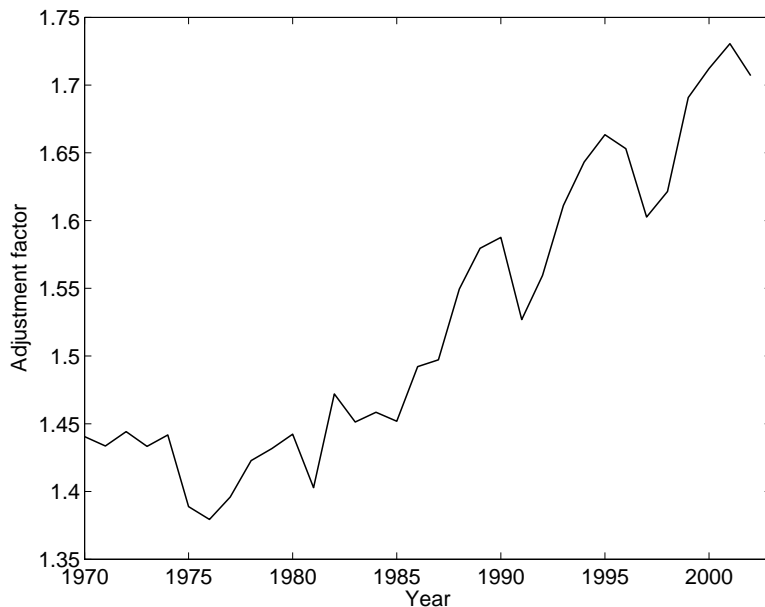
Figure 11: The adjustment factor $a_t/(a_t - 1)$ that topcoded earnings are multiplied by under the assumption that the top decile of the earnings distribution is Pareto distributed.

adjustment factor for our topcoded earnings. We obtain year-specific 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 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 11 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 12 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 10. 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.
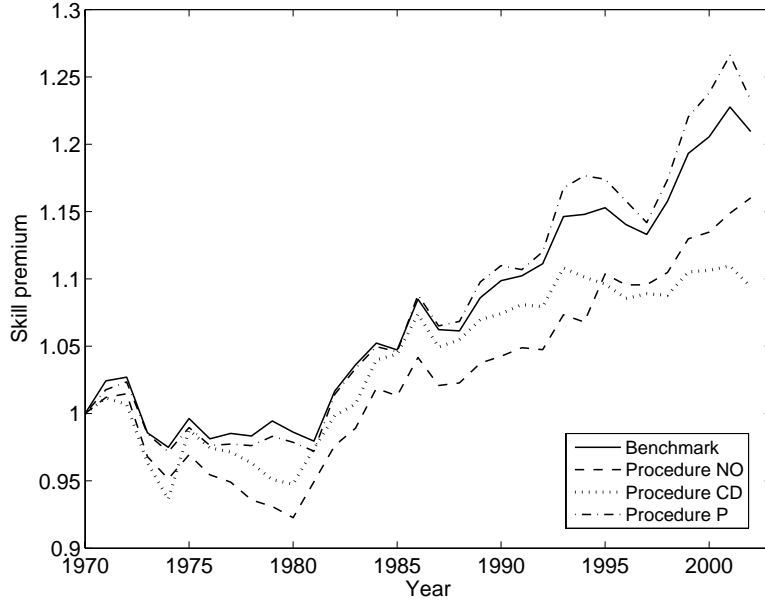
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Figure 12: Skill premium in United States for various procedures to handle topcoded earners.

Appendix A.5   Income shares in Sweden

The income share for labor (χ) is calculated as the annual average over the period 1970-2002. For the period after 1994 we have used data from Statistics Sweden's web site www.scb.se, while earlier data is published in Statistical reports N1981:2:5 (Appendix 4, "Production and factor income 1963-1980") and N 10 SM 9501 ("National accounts 1980-1994"). We compute a labor share of income equal to 0.61 in the private sector and 0.68 in the total economy.

The income share for structures (α) is calculated as the annual average over the period 1970-1995, using stocks of structures and equipment and assuming that the rental rates on structures and equipment are given by $r_s = 0.09$, and $\tilde{r}_e = 0.209$, respectively, as implied by premises a)-c) in section 4.2. There exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon (1990), but noting that $r_e k_{e,t} = \tilde{r}_e \hat{k}_{e,t}\, \phi$ where $\hat{k}_{e,t}$ is the unadjusted stock of equipment at time t, $\phi = \left(1 - \frac{(1-\delta_e)}{1+\triangle}\right) / \left(1 - \frac{(1-\delta_e)p'_e/p_e}{1+\triangle}\right)$, and $\triangle$ is the average growth rate of investment in unadjusted equipment, we calculate the income share of structures as

$$\alpha = \frac{(1-\chi)}{26} \sum_{t=1970}^{1995} \frac{r_s k_{s,t}}{r_s k_{s,t} + \tilde{r}_e \hat{k}_{e,t}\, \phi}\,.$$

The data on equipment and structures are from Statistics Sweden Statistical reports N1981:2:5 (Appendix 2, "Capital formation and stocks of fixed capital 1963-1980") and N 10 SM 9501 (Appendix 3, "Stocks of fixed assets and national wealth 1980-1995"). The average growth rate of investment in equipment is 3.5 percent. We compute an income share of structures

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equal to 0.17 in both the private sector and in the total economy.

Appendix B KORV's Model

Appendix B.1 Equilibrium properties

By solving for k_s from first-order condition (3) and substituting into production function (2), we get an equilibrium expression for output

$$y = (A\,\Omega)^{\frac{1}{1-\alpha}} \left[\frac{\alpha}{r_s}\right]^{\frac{\alpha}{1-\alpha}} \left\{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}}\right\}^{\frac{1}{\sigma}}. \tag{10}$$

From first-order conditions (4) and (6), we get an equilibrium expression for the quantity of equipment,

$$k_e = \left[\frac{w_s}{\theta_s r_e}\right]^{\frac{1}{1-\rho}} h_s. \tag{11}$$

From first-order conditions (5) and (6), we get an equilibrium expression for the labor share of income, χ,

$$\chi \equiv \frac{w_u h_u + w_s h_s}{A\,F(k_s, k_e, h_u, h_s)} = (1-\alpha)\frac{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}}\theta_s h_s^\rho}{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}}}$$

$$= (1-\alpha)\frac{1 + \pi\,\dfrac{h_s}{h_u}}{1 + \left[\pi^\sigma\,\theta_u^\rho\,\theta_s^{-\sigma}\left(\dfrac{h_s}{h_u}\right)^{\sigma(1-\rho)}\right]^{\frac{1}{\sigma-\rho}}}, \tag{12}$$

where $\pi \equiv w_s/w_u$ is the skill premium, and the last equality in (12) is obtained by twice invoking expression (9).

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 income, 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 (10) and (11) for output and equipment, we confirm that first-order condition (6) 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 (5) 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 income in (12) 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 to the base year

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

1. For an arbitrary value of Ω, we can compute the labor share of income as a function of the parameters θ_u and θ_s. Such a mapping is illustrated in figure 13 for one particular value of Ω. Given a calibration value for the labor share, say $\chi = 2/3$, the bold curve in figure 13 depicts pairs of parameter values (θ_u, θ_s) that produce the target value χ.

2. Both the labor share of income and the skill premium are constant along the bold curve in figure 13, as described in Claim 1. Hence, given the calibration value of the labor share, the bold curve in figure 13 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 14.

3. Given a calibration value for the skill premium, figure 14 pins down our choice of parameter Ω. Given that value of Ω, the associated figure 13 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.
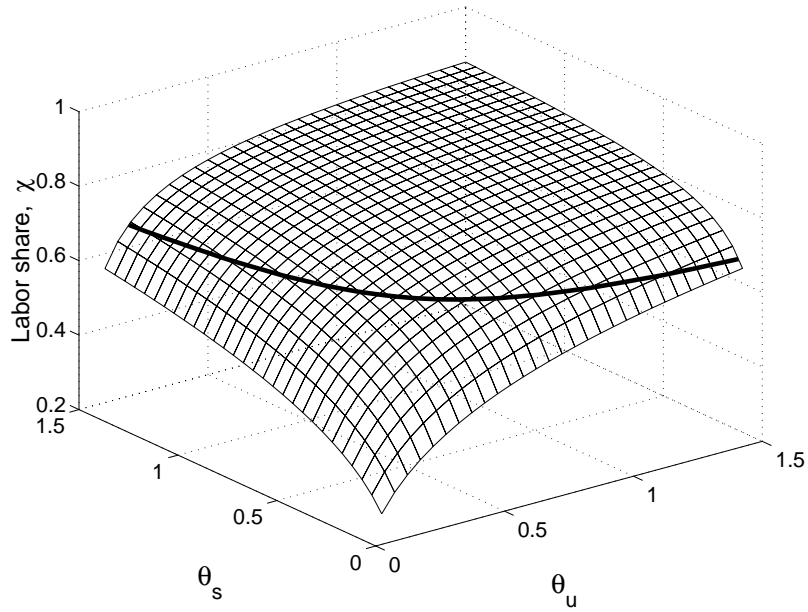
Figure 13: Labor share of income, χ, as a function of parameters θ_u and θ_s, for a given value of the parameter Ω. Along the bold curve is not only the labor share constant, $\chi = 2/3$, but so is the skill premium.
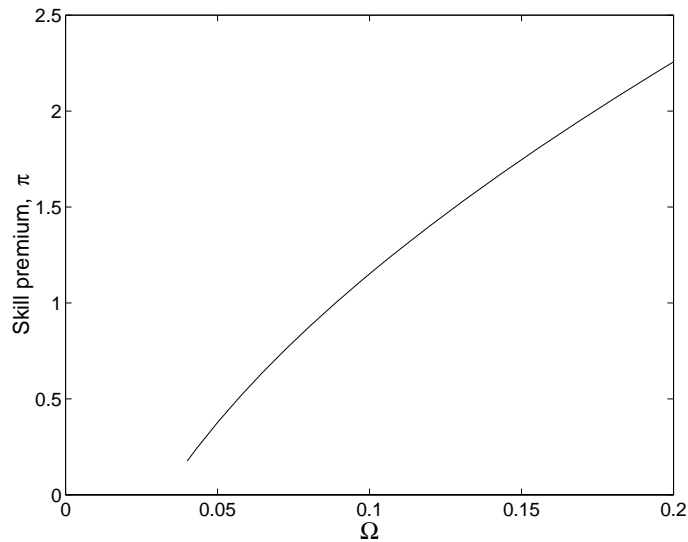

Figure 14: Skill premium, π, as a function of the parameter Ω, given that the labor share of income is held constant, $\chi = 2/3$.