The Missing Swedish Skill Premium:
Sweden versus the United States 1970-2002

David Domeij and Lars Ljungqvist*

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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 service the welfare state.

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


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*Domeij: Stockholm School of Economics (email: david.domeij@hhs.se); Ljungqvist: Stockholm School of Economics and New York University (email: lars.ljungqvist@hhs.se). We have received helpful comments at the 2005 SED conference and from numerous seminar participants. Financial support from The Jan Wallander and Tom Hedelius Foundation is gratefully acknowledged.
1 Introduction

“Sweden defies economic theory” is not a seldomly heard phrase. One major puzzle is the apparently meager reward to skills in the Swedish labor market. This paper sets out to provide a resolution to this puzzle by first, computing as accurately as possible a measure of the Swedish skill premium based on population data and second, using a structural model to rationalize the skill premia of Sweden and the U.S. 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 collapse of the Swedish skill premium which has been truly spectacular. 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. This has led many observers to propose nonmarket explanations to the evolution 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.\(^1\) In contrast, our goal is to 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.\(^2\) The historical increase in the relative supply of skilled labor is more than offset by skill-biased technological change, as measured 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.

\(^1\)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.”

\(^2\)For another study that explains the U.S. skill premium in terms of observed labor quantities, see Katz and Murphy (1992). They estimate 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 shift for skilled labor. Edin and Holmlund (1995) employ a version of that regression analysis on Swedish data.
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.\(^3\) 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 market forces equalize the wage of each type of labor across sectors. And if perfect competition prevails in the private sector, labor is paid its marginal product in private employment. It follows that, under the assumption of constant returns to scale, 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.

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.\(^4\) 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 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 from the private sector into public sector employment, long-term unemployment, or disability and early retirement programs.

The rest of the paper is organized as follows. Section 2 describes the data, and section 3 lays out KORV's model. The skill premia in the U.S. and Sweden are simulated in section 4. Section 5 concludes with a discussion of our findings. Two appendices provide some further details on the data and the theoretical 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 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.

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\(^{3}\) 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. 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 overall economy. Total output is smaller than it would have been if household services had been paid for privately and transacted through the market." (Rosen, 1997, pp. 104–105.)

\(^{4}\) The predictive power of the model for the U.S. is largely unaffected by the choice of either private or total employment.
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 within-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.\(^5\) 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.\(^6\)

2.3 Trends in skill premia and labor supply

We lay out the basic facts in figures 1 through 3. Figure 1 shows how the Swedish and U.S. skill premia evolved between 1970 and 2002.\(^7\) The U.S. development is well known. The

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\(^5\) See Appendix A.4 for a discussion of topcoding in the U.S. data.

\(^6\) 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.

\(^7\) 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.
skill premium deteriorated during the 1970s, but has grown rapidly since around 1980. 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 2 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. Figure 3 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 college educated workers

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8Our 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.

9We 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.
in Swedish private employment has grown much more than that in total employment in figure 2. As we describe in section 5.1, 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 2 and 3 as given, and then explain the skill premium series in figure 1.

3 The Model

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

\[ y = A k_s^{\alpha} \left\{ \mu (\psi, h_u)^\sigma + (1 - \mu) \left[ \lambda h_c^\sigma + (1 - \lambda) (\psi, h_s)^\sigma \right] \right\}^{\frac{1}{\sigma}}, \]

where \( y \) is aggregate output, \( k_s \) and \( k_c \) 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 \( \alpha \), and \( \sigma \) and \( \rho \) determine elasticities of substitution. 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 \( 1/(1 - \sigma) \).\(^{10}\)

Income shares of

\(^{10}\)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
Figure 3: 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).

different production factors are governed by $\mu$ and $\lambda$, and $\psi_u$ and $\psi_s$ convert hours of unskilled and skilled labor, respectively, into efficiency units.

Rates of depreciation on structures and equipment are given by $\delta_s$ and $\delta_e$, 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,

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

where $\Omega \equiv (1 - \mu)^{\frac{1-\alpha}{2}} \lambda^{\frac{1-\alpha}{2}}, \theta_u \equiv \mu \psi_u^\sigma (1 - \mu)^{-1}\lambda^{-\frac{\sigma}{2}}$, and $\theta_s \equiv (1 - \lambda)\psi_s^\sigma \lambda^{-1}$. 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.
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{align}
  r_s &= A \alpha F(k_s, k_e, h_u, h_s) k_s^{-1}, \\
  r_e &= A \Gamma(k_s, k_e, h_u, h_s) (k_e^p + \theta_e h_e^p)^\frac{\alpha}{\sigma} k_e^p^{-1}, \\
  w_u &= A \Gamma(k_s, k_e, h_u, h_s) \theta_u h_u^{\sigma-1}, \\
  w_s &= A \Gamma(k_s, k_e, h_u, h_s) (k_e^p + \theta_e h_e^p)^\frac{\alpha}{\sigma} \theta_e h_e^{\sigma-1},
\end{align}

respectively, where

\[ \Gamma(k_s, k_e, h_u, h_s) \equiv (1 - \alpha) \frac{F(k_s, k_e, h_u, h_s)}{\theta_u h_u^{\sigma} + (k_e^p + \theta_e h_e^p)^\frac{\alpha}{\sigma}}. \]

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.\(^{11}\) Hence, the rental rates satisfy

\[ r_s = r + \delta_s, \]

\[ \frac{r_e}{p_e} = 1 + r - \frac{\delta_e}{p_e} (1 - \delta_e) \equiv \tilde{r}_e, \]

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,

\[ \frac{w_s}{w_u} = (k_e^p + \theta_e h_e^p)^\frac{\alpha}{\sigma} \frac{\theta_e h_e^{\sigma-1}}{\theta_u h_u^{\sigma-1}}. \]

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

### 3.1 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 4.3.) Next, given

\(^{11}\)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.
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 \((\delta_s, \delta_c)\), and equipment price movements \((p'_e/p_e)\), which are assumed to be the same across countries. We use country-specific historical averages of factor income shares from the national income accounts, including the capital structures share of income \((\alpha)\).

To complete the parameterization of the U.S. model, we use KORV’s estimates for the elasticity parameters \(\sigma\) and \(\rho\). In the absence of comparative Swedish estimates, we need to proceed differently for Sweden.\(^{12}\) We split the Swedish data in two subperiods: 1970–1990 and 1990–2002. We use data from one subperiod to calibrate the elasticities, and data from the other subperiod to assess the model’s predictive power. 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.

For further details about the calibration procedure, see Appendix B.2.\(^{13}\)

### 3.2 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.

\(\text{a)} \) The real interest rate is set equal to \(r = 0.04\), which is 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.

\(\text{b)} \) 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

\(^{12}\)Lindquist (2005) has recently estimated KORV’s model using Swedish data, but his proxy of the skill premium is a poor one. According to Lindquist, the Swedish skill premium in 1999 is back to its 1970 value whereas we showed, using census data and tax records, that it is some 20 percent below.

\(^{13}\)Appendix B.2 describes how the remaining scale parameters, \(\{\Omega, \theta_u, \theta_s\}\), are calibrated based on outcomes in the base year. But, of course, data from other years enter indirectly since the calibration uses historical averages of rates of return and factor income shares, and estimates of technology parameters, all of which reflect economic outcomes over longer periods of time.
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, \( \hat{r}_e = 0.209 \), 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 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.

4.1 U.S. skill premium

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 (\( 1/(1 - \rho) \)).

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 4 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. All other inputs in the calibration are taken off the shelf from other studies. 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.
Figure 4: 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.

4.1.1 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.

4.2 Swedish skill premium

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.\textsuperscript{14} We now turn to our method for calibrating the Swedish elasticities of substitution.\textsuperscript{15}

4.2.1 Elasticities of substitution

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 ($\sigma$, $\rho$). In particular, for each $\rho$ that belongs to a set that we have picked, we compute the value of $\sigma$ that best captures the observed skill premium in the calibration period (using the least-squares metric). Given each such pair ($\sigma$, $\rho$), we then simulate the model over the whole period 1970–2002.

We pick the set of $\rho$ 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 for giving 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).

\textsuperscript{14}For details on the computations of factor income shares, see Appendix A.5. We want to emphasize that the ensuing analysis is quite robust to the calibration of factor income shares. 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.

\textsuperscript{15}Our motivation for considering different elasticities in Sweden and the U.S. are twofold. On the empirical side, Bergström and Panas (1992) estimate relatively low elasticities of substitution between unskilled and skilled labor in many Swedish industries. On the theoretical side, we conjecture that the much lower 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. Specifically, using our weighting scheme for computing quantities of labor inputs, the skilled fraction of the labor force in private employment in 1970 (2002) constituted 3.6% (16.4%) in Sweden as compared to 16.4% (36.6%) in the U.S.
Figure 5: Swedish skill premium, given calibration period 1970-90, in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

4.2.2 Private versus total employment

The upper panel of figure 5 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 5, 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 6 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 interval.  

In figure 7, we use the whole period 1970–2002 to calibrate the substitution elasticities. As compared to figure 5, the upper panel based on private employment in figure 7 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 7, the path of the simulated skill premium can be understood in terms of the above out-of-sample calibration-simulation

\[ (1 - \rho)^{-1} \in [0.3, 0.8]. \] 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 correspond 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.
Figure 7: 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.

procedures. As shown in the lower panels of figures 5 and 6, 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 5 and 7 are qualitatively the same – the 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.\textsuperscript{17}

\textsuperscript{17}The base year in figure 7 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.
4.3 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 7. Under the auxiliary premise 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.

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.

<table>
<thead>
<tr>
<th>Table 1: Model simulations 1970–2002</th>
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<tr>
<td>Average labor share of income</td>
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<tr>
<td>(calibrated value in 1970)</td>
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<tr>
<td>Annual productivity gain(^a)</td>
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<tr>
<td>Annual growth rate of equipment</td>
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\(^a\) Measured ‘per hour,’ i.e., we normalize \( h_u + h_s = 1 \) in each year.

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). 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.

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

\(^{18}\) We have picked the calibration of the Swedish model with elasticity \((1 - \rho)^{-1} = 0.35\) and the associated choice of \( \sigma \) that yields the best fit of the simulated skill premium, \((1 - \sigma)^{-1} = 1.21\).

\(^{19}\) 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.

\(^{20}\) Using data from the Groningen Growth and Development Center (http://www.ggdc.net, the Total Economy Database, te07l.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.
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. While not conclusive, this suggests that the stocks of equipment developed similarly in both countries.

5 Discussion

5.1 Public sector employment

We are not the first ones to suggest that public sector employment can be an important determinant for wage outcomes. 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” labor. As we have argued above and as suggested 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 sector. Recall figure 3 that depicts the ratio of skilled/unskilled labor in private employment 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

---


22Edin 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.)
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.

5.2 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’ compensation are competitively determined.

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 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 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 economics’ skill formation and how such skill formation responds to changing rates of return to education as implied by a changing skill premium.23 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.

23It 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.
5.3 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 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 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.\(^{24}\) A hypothesis that merits further investigation.

\(^{24}\)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 northeast 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.
References


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 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.\textsuperscript{25} 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.\textsuperscript{26} 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.

\section*{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

\textsuperscript{25}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 wage are less than 40 times one sixth of the hourly minimum wage.

\textsuperscript{26}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.
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 $h_{jt} = \sum_{g \in G_j} H_{gt} \tilde{W}_g v_{gt}$, where the weights $\tilde{W}_g$ are the average group wages for 1970 and 1990, 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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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
\[
r_{it}^{hrs} = \frac{e_{it}}{W_{it} r_{it}^{wks}}, \quad \text{for } i \in g \text{ and } d_{it} = 0.
\]

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 within 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.\(^{28}\) 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 Poljolassa, 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 men and underestimates the working hours of women. To investigate the importance of this mismatch we perform the following robustness check. The midpoints in each interval are multiplied by two factors; the first capturing aggregate time affects and the 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 worked 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.

Appendix A.4 Topcoding in U.S. data

As mentioned above, data on wage and salary income is topcoded in U.S. data. Figure 8 shows how the share of individuals who are topcoded in our wage sample has increased over time (but whenever the top-coded 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 conditions.

\(^{28}\) 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.
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,\footnote{For a recent study using the Pareto distribution to interpolate income distributions, see Piketty and Saez (2003).} 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-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., $\vartheta_t(e) = (b/e)^{a_t}$, 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 $\{\vartheta_t(e_y), e_y\}$ are constructed from the earners in the top

Figure 8: 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.
Figure 9: 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.

decile who are not topcoded. (The \( R^2 \) values in these regressions are above 0.98.) Figure 9 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 10 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 8. 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 Income shares in Sweden

The income share for labor (\( \chi \)) 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.
Figure 10: Skill premium in United States for various procedures to handle topcoded earners.

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 \( \hat{r}_e = 0.209 \), respectively, as implied by premises a)-c) in section 3.2. There exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon (1990), but noting that \( r_e k_{e,t} = \hat{r}_e \hat{k}_{e,t} \phi \) where \( \hat{k}_{e,t} \) is the unadjusted stock of equipment at time \( t \), \( \phi = \left(1 - \frac{1-\delta_e}{1+\Delta}\right) / \left(1 - \frac{1-\delta_e}{1+\Delta}\right) \), and \( \Delta \) is the average growth rate of investment in unadjusted equipment, we calculate the income share of structures as

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

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 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

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

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

$$k_s = \left[ \frac{w_s}{\theta_s r_s} \right]^{\frac{1}{1-\alpha}} h_s. \tag{11}$$

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

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

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

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 $\{r_s, r_c\}$, suppose there exists an equilibrium $\{w_s, w_u = w_s/\pi, \chi, k_s, k_c\}$. Then by varying the initial parameters $\{\theta_u, \theta_s\}$, there exists a continuum of other equilibria, $\{\hat{w}_s, \hat{w}_u = \hat{w}_s/\pi, \chi, \hat{k}_s, \hat{k}_c\}$, where the skill premium, $\pi$, and the labor share of income, $\chi$, 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 \( \chi \) and skill premium \( \pi \) in the base year, while allowing for any absolute wage level \( w_s \).

1. For an arbitrary value of \( \Omega \), we can compute the labor share of income as a function of the parameters \( \theta_u \) and \( \theta_s \). Such a mapping is illustrated in figure 11 for one particular value of \( \Omega \). Given a calibration value for the labor share, say \( \chi = 2/3 \), the bold curve in figure 11 depicts pairs of parameter values \( (\theta_u, \theta_s) \) that produce the target value \( \chi \).

2. Both the labor share of income and the skill premium are constant along the bold curve in figure 11, as described in Claim 1. Hence, given the calibration value of the labor share, the bold curve in figure 11 represents a mapping from the parameter \( \Omega \) (that is held fixed in the graph) to a specific value of the skill premium, \( \pi \). Given the calibration value of the labor share, the compilation of pairs \( (\Omega, \pi) \) from successive graphs when varying \( \Omega \), yields a mapping as illustrated in figure 12.

3. Given a calibration value for the skill premium, figure 12 pins down our choice of parameter \( \Omega \). Given that value of \( \Omega \), the associated figure 11 contains a bold curve depicting permissible parameter values \( \theta_u \) and \( \theta_s \). On the basis of Claim 1, we can find parameter values \( (\theta_u, \theta_s) \) to match any absolute wage level; we set \( w_s = 1 \) for both countries.
Figure 11: Labor share of income, $\chi$, as a function of parameters $\theta_u$ and $\theta_s$, for a given value of the parameter $\Omega$. Along the bold curve is not only the labor share constant, $\chi = 2/3$, but so is the skill premium.

Figure 12: Skill premium, $\pi$, as a function of the parameter $\Omega$, given that the labor share of income is held constant, $\chi = 2/3$. 