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Abstract

The last three decades have seen a sharp increase in the skill premium for college educated in the United States while workers in Europe have experienced a compression of wage structures with lower skill premia, which is perhaps starkest seen in Sweden. The U.S. skill premium rose by around 21 percent between 1970 and 2002. In Sweden, the skill premium fell by more than 30 percent(!) between 1970 and 1990 and then recovered by less than 10 percentage points in 2002. These diametrically different outcomes pose a challenge to economic analysis based on times series of aggregate labor supplies. Under the assumption that Sweden has faced the same skill-biased technological change as estimated for the US, we find that Swedish relative supplies of skilled and unskilled labor cannot explain the observed dramatic decline in the skill premium.

In our solution to this puzzle, we first emphasize that competitive equilibrium theory predicts that the skill premium is determined by marginal products in the private sector. Second, we question the common approach of computing these marginal products by using economy-wide labor supplies, since it presupposes that public-sector employees perform tasks that are perfect substitutes to those performed by private-sector employees. Under the alternative assumption that public sector employment does not affect marginal products in the private sector, we can explain the disparate developments of skill premia in the US and Sweden. The dramatic decline of the Swedish skill premium 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 system. 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.
1 Introduction

The last three decades have seen a sharp increase in the skill premium for college educated in the United States while workers in Europe have experienced a compression of wage structures with lower skill premia, which is perhaps starkest seen in Sweden. According to Table 1, the US skill premium rose by around 9 percent between 1970 and 1990 and an additional 12 percentage points in 2002. During the same period 1970-1990, the Swedish skill premium fell by more than 30 percent(!) and recovered by less than 10 percentage points in 2002. These numbers are computed by dividing all employees into “skilled” and “unskilled” workers where skilled are those with a traditional college education (at least three years of university studies in Sweden and a total of at least 16 years of schooling in the U.S.). The Swedish data is taken from census data and tax records which include Sweden’s entire population, while the American data is drawn from U.S. Department of Commerce’s Current Population Survey (CPS). As described in the data appendix, we use 1970 as a base year which implies that the wage ratio between skilled and unskilled labor is normalized to 1 in 1970 and hence, we focus on the relative change in skill premium in 1990.

Table 1: Skill premium in the U.S. and in Sweden
(relative wage of skilled versus unskilled labor)

<table>
<thead>
<tr>
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<th>USA</th>
<th>Sweden</th>
</tr>
</thead>
<tbody>
<tr>
<td>Skill premium</td>
<td>1.099 1.210</td>
<td>0.687 0.790</td>
</tr>
</tbody>
</table>

These diametrically different outcomes are not a result of fundamentally different initial educational wage distributions; in 1970 the U.S and Sweden were rather similar. The absolute level of relative wages between an engineer and a blue-collar worker were 1.6 in the U.S. and 1.7 in Sweden.1 This has led many observers to resort to non-market outcomes when explaining the evolution of the Swedish skill premium. For example, Edin and Topel (1997) suggest that the answer should be sought in a conspiracy theory involving the trade union and the employers’ association.2 Moreover, as Sweden is a small open economy, it is difficult to argue that Sweden has not been exposed to similar skill-biased technological changes and forces of globalization as the U.S.

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

1 See the data appendix for a detailed description of these calculations.
2 See also Hibbs (1990) and Davis and Henrekson (2005).
We first show that using total Swedish relative supplies of skilled and unskilled labor cannot explain the observed dramatic decline in the skill premium but rather, the analysis predicts counterfactually an increasing premium. We then show that under two important qualifications, relative supplies of skilled and unskilled labor can explain the observed dramatic shifts. The qualifications are the following.

We first emphasize that competitive equilibrium theory predicts that the skill premium is determined by marginal products in the private sector. Second, we question the common approach of computing these marginal products by using economy-wide labor supplies, since it presupposes that public-sector employees perform tasks that are perfect substitutes to those performed by private-sector employees. Hence, the proper objects of our study are not economy-wide labor aggregates but rather labor inputs in the private sector. This distinction turns out to be immaterial for the study of the U.S. skill premium but of critical importance when explaining Swedish labor market outcomes.

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

The expansion of the public sector does not have a direct effect on the skill premium, as any economic model with free labor mobility implies that changes in the skill premium should be the same across sectors—a implication that is also born out in the data as we show below. Instead, the influence of the public sector upon the skill premium is an indirect one where public employment decisions affect the ratio of skilled to unskilled workers available to the private sector. The numbers and skill composition of public employees are not constrained by profit considerations in the market place, but rather guided by public policies and largely financed through taxation. Indeed, the expansion of the Swedish public sector has disproportionately involved the hiring of unskilled workers. According to Table 2, in the U.S. the logarithmic change in ratio of skilled to unskilled was of similar magnitude in private sector employment and in total employment. In Sweden on the other hand, the ratio of skilled to unskilled workers has grown more rapidly in the private sector as compared to the overall Swedish economy.

\[3 \text{In fact, economic models with free labor mobility imply that wages are equalized across sectors for the same type of labor. In our study, we do not address the common assertion of seemingly lower levels of earnings in the public sector as compared to the private sector.}\]
In the paper we consider two different approaches in analyzing the relationship between changes in relative labor supplies and the evolution of the skill premium. First, we follow Katz and Murphy (1992) and assume that relative demand effects can be captured using a simple linear trend. Second, we follow Krusell, Ohanian, Ríos-Rull and Violante (2000) and assume that relative demand effects are due to capital-skilled complimentarity in combination with equipment-specific technological progress.

Under the alternative assumption that public sector employment does not affect marginal products in the private sector, we can explain the disparate developments of skill premia in the US and Sweden. 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 system.

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 paper is organized as follows. In Section 2 the data is described and the facts are presented. Section 3 contains the analysis using the approach in Katz and Murphy (1992) and ?? contains the analysis using approach in Krusell et al (2000). Section 6 concludes with a discussion of the results.

2 Data and Facts

In this section we briefly describe the data and how we construct our measures for skill premium and labor input. 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 Appendix A.

2.1 U.S. data

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

<table>
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<tr>
<th></th>
<th>USA</th>
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<th>Sweden</th>
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<tbody>
<tr>
<td>Private sector</td>
<td>0.82</td>
<td>1.08</td>
<td>1.24</td>
<td>1.68</td>
</tr>
<tr>
<td>Aggregate economy</td>
<td>0.67</td>
<td>0.90</td>
<td>0.85</td>
<td>1.22</td>
</tr>
</tbody>
</table>

Table 2: Ratio of skilled to unskilled labor in the U.S. and in Sweden
into groups according to age, sex, race, education, and private/public sector employment.

We follow Katz and Murphy (1992) and 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. For the aggregation we follow Krusell et al (2000). Total labor supply for each class is given by weighting hours in each group within the class by average group wages for 1970-2002 and then summing. The average hourly wage for each class 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 as on our U.S. 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 covers 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 data source 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, and since all individuals are included there are no problems associated with topcoding.

For these years we thus obtain as close as possible to a perfect measure of the skill premia and of the relative supply of skilled and unskilled labor. When discussing our results, emphasis will therefore be placed on these observations.

For the years 1971 through 1989 there exist no comparable data source in Sweden. To trace out the development of the skill premia for these years we use results from a study by Edin and Holmlund (1995). Using survey data they estimate human-capital type wage equation for the 1970's and 1980's. We assume that our measure of the skill premium has the same time profile as Edin and Holmlund's estimated wage differential between workers with 16 and 12 years of education (see Table 9.2, p. 318). It is comforting to note that the change in their estimated wage differential between 1968 and 1991 is very similar to the change in the skill premia we derive using the Census data from 1970 and 1990.

To trace out the development of labor input for the years 1971 through 1989, we rely on data from labor-market surveys (Arbetskraftsundersökning, AKU) that also gather information on workers’ educational attainment and industry classification. For these years we assume that our measure of labor input has the same time profile as the ratio of number of skilled to unskilled in the survey.

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4 This differs from Katz and Murphy (1992) who aggregate using a wage index constructed from average employment shares.
5 See appendix Appendix A for a discussion of topcoding in U.S. data.
7 We thank Anders Edin for providing us with the data. Individuals with SNI code 91,931-934 are classified as public sector employees.
data. It is once more comforting to note that the change in the ratio of skilled to unskilled in the survey data over the period 1971 and 1990 is very similar to the change in our measure of labor input derived from the Census data.

2.3 Trends in labor supply and skill premia

We lay out the basic facts in Figures 1 and 2. Figure 1 shows a substantial rise in the relative supply of college educated workers in both Sweden and the U.S. The supply of skilled workers rose by 200 percent between 1970 and 2002 in the U.S. The percentage change is even larger in Sweden. Between 1970 and 2002 the supply of skilled workers rose by 540 percent.

Figure 2 shows how the Swedish and U.S. skill premia evolved between 1970 and 2002. The U.S. development is well-known. The skill premia deteriorated during the 1970’s, but has grown rapidly since around 1980. The development of the Swedish skill premia is dramatically different. Although the period it remained below its 1970 value, and by 1990, the skill premia had fallen by an astonishing 31 percent. Thereafter it recovered somewhat, but in 2002 it was still 20 percent smaller than in 1970.

3 A relative wage equation

Katz and Autor (1999), in their chapter in the Handbook of Labor Economics, review a common approach of studying changes in educational wage differentials

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6 This picture is very similar to that documented by Autor et al (2005). Recently, some authors have argued that the rise in skill-premia during the 1980’s was an episodic event rather than part of a secular trend. As shown in the appendix, however, this argument relies to large extent on the chosen topcoding procedure.
by breaking the work force into two broad educational groups, say college equivalents and high school equivalents ("skilled" and "unskilled" workers). Katz and Autor illustrate the approach by considering a production function with a CES specification over labor inputs, which we here nest inside a Cobb-Douglas specification with capital,

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

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

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

\[ \pi \equiv \frac{w_s}{w_u} = \frac{1 - \mu}{\mu} \left( \frac{\psi_s}{\psi_u} \right)^\sigma \left( \frac{h_s}{h_u} \right)^{(1-\sigma)} \]  

Katz and Murphy (1992) provide a successful empirical implementation of this approach to explain changes in the U.S. college/high school wage differential from 1963 to 1987. They assume that a simple linear time trend can approximate the impact of changing parameters \( \{\mu, \psi_u, \psi_s\} \) upon the logarithm of the wage differential. After computing the logarithmic change of the skill premium in expression (2) and using Katz and Murphy’s approximation with their estimated coefficients, we arrive at the following candidate explanation to the skill
Figure 3: Actual and predicted skill premia. The solid lines refer to actual skill premia and the dashed lines depict predicted skill premia.

premium development,

\[
\log (\pi_t) = \log \left( \frac{1 - \mu_t}{\mu_t} \right) + \sigma \log \left( \frac{\psi_{s,t}}{\psi_{u,t}} \right) - (1 - \sigma) \log \left( \frac{h_{s,t}}{h_{u,t}} \right) \\
\approx \text{constant } + 0.033 \times \text{time} - 0.709 \log \left( \frac{h_{s,t}}{h_{u,t}} \right), \tag{3}
\]

where Katz and Murphy (1992) estimate that the time trend increases the skill premium by 3.3 log points annually and the elasticity of substitution between skilled and unskilled labor is equal to \(1/(1 - \sigma) = 1/0.709 \approx 1.41\). As noted by Katz and Autor (1999), the estimated elasticity is in the middle of the range of 0.5–2.5 in earlier studies using cross-sectional approaches reviewed by Freeman (1986).

Figure 3 compares the actual and predicted skill premia for the period 1970-2002 where the predicted skill premia are derived by substituting the ratio of skilled to unskilled workers in the private sector from Figure 1 into the Katz-Murphy model; equation (3). The Katz-Murphy model does a very good job in explaining the U.S. skill premia up til the early 1990’s, but predicts a more rapid growth in the skill premium than occurred in the last decade. This has previously been noted by Katz and Autor (1999), Card and DiNardo (2002), and Autor et al (2005), who interpret this as a slowdown in the relative demand for skilled workers. For example, Autor et al reestimate equation (3) using data from 1963-2003 and allowing for trend break in 1992. Their results indicate a significant slow-down of demand after 1992, but they also reconfirm the importance of the relative supply growth.

Concerning the Swedish skill premium, the Katz-Murphy model predicts a very similar time path as that observed in the data. But, even though the Katz-Murphy model predicts a large drop, it falls short of the even more spectacular observed decline.

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3.1 Earlier study using the framework

This framework has previously been applied on Swedish data by Edin and Holmlund (1995). They also find that the Katz-Murphy model goes a long way in explaining the observed Swedish skill premia. However, in contrast to ours, their finding requires that (i) the increase in the relative demand of skilled workers was much smaller in Sweden than in the U.S., and (ii) that the substitutability between skilled and unskilled in the production is much large in Sweden than in the U.S.. The reason turns out be due to the choice of data. In their benchmark specification Edin and Holmlund use the university wage premium among male white-collar workers in manufacturing, mining and construction (uncorrected for age differences) as a proxy for the economy-wide university wage premiums. It turns out to be a poor proxy. Between 1970 and 1990, it only fell by some 10 percentage points, whereas we showed that the economy-wide skill premia fell by more than 30 percent. As a result, when estimating equation (3) Edin and Holmlund find (i) small relative demand effects as captured by the time trend and (ii) an elasticity between university and high school graduates of 2.9 which is much larger than those reported for other countries by Katz and Autor (1999). In a sensitivity analysis they also consider the economy-wide university wage premium, but without including the relative demand effect (the time trend). Again their estimated elasticity is very large; around 2.5.

It is therefore interesting to note that estimating equation (3) on our Swedish data for the 'same' period (1971-90) yields estimates remarkably similar to those of Katz and Murphy;

\[
\log (\pi_t) = 0.030 \times \text{time} - 0.744 \log \left( \frac{h_{s,t}}{h_{u,t}} \right) + \text{constant}, \quad \hat{R}^2 = 0.85, \quad (4) \tag{2.27} (3.68)
\]

where the numbers in parentheses are absolute t-values.

4 A model with capital-skill complementarity

We now analyze the skill premium in terms of the production function proposed by Krusell, Ohanian, Ríos-Rull and Violante (2000), hereafter referred to as KORV. They distinguish between capital structures, \(k_s\), and capital equipment, \(k_c\), by augmenting production technology (1) as follows,

\[
y = AG(k_s, k_c, \psi_u h_u, \psi_s h_s)
\]

where the function \(G\) is is given by

\[
G(k_s, k_c, \psi_u h_u, \psi_s h_s) \\
\equiv k_s^\alpha \left\{ \mu (\psi_u h_u)^\rho + (1 - \mu) [\lambda k_c^\rho + (1 - \lambda) (\psi_s h_s)^\rho]^{1/\rho} \right\}^{1/\alpha}.
\]

The new parameters are \(\lambda\), governing income shares, and \(\rho\), determining substitution possibilities. 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 $1/(1 - \sigma)$.

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 particular, one unit of output can be converted into $1/p_t$ units of capital equipment at time $t$. In a competitive equilibrium, $p_t$ becomes the relative price of capital equipment.

Since the production technology exhibits constant returns to scale, we can normalize the sum of skilled and unskilled labor hours to one. From hereon, let $h_s$ denote the skilled labor share of total hours, at which production function (5) becomes constant over time for capital structures, they introduce equipment-technology for converting output into new capital but while the transformation can be written as

\[ y = AG(k_s, k_e, \psi_u(1 - h_s), \psi_s h_s) \]

\[ = A \Omega k_s^{\alpha} \left\{ \theta_u (1 - h_s)^\sigma + (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{1}{\rho}} \right\}^{\frac{1}{1-\rho}} \]

\[ \equiv AF(k_s, k_e, h_s), \quad (6) \]

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

KORV mention that any one of the four parameters $\{\mu, \lambda, \psi_u, \psi_s\}$ in production function (5) must be fixed as a normalization in their estimation. This is made explicit in our production function (6) where the four parameters are mapped into three, $\{\Omega, \theta_u, \theta_s\}$.

### 4.1 Equilibrium characterization

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

\[ r_s = A \alpha F(k_s, k_e, h_s) k_s^{-1}, \quad (7) \]

\[ r_e = A \Gamma(k_s, k_e, h_s) (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{1}{\rho}} k_e^{\rho-1}, \quad (8) \]

\[ w_u = A \Gamma(k_s, k_e, h_s) \theta_u (1 - h_s)^{\sigma-1}, \quad (9) \]

\[ w_s = A \Gamma(k_s, k_e, h_s) (k_e^{\rho} + \theta_s h_s^{\rho})^{\frac{1}{\rho}} \theta_s h_s^{\rho-1}, \quad (10) \]

respectively, where

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

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.\(^\dagger\)

\(^\dagger\)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.

\(^\ddagger\)Following KORV we ignore risk premia in our analysis of capital investments. Furthermore, to simplify the notation, our expressions are written under perfect foresight.
Hence, the rental rates satisfy
\[ r_s = r + \delta_s, \]
\[ r_e = p_t \left[ 1 + r - \frac{p_t + 1}{p_t} (1 - \delta_e) \right] \equiv p_t R_e, \]
where \( \delta_s \) and \( \delta_e \) are the rates of depreciation on structures and equipment. The rental rate \( r_e \) is equal to the marginal product of one unit of equipment. Since \( 1/p_t \) units of equipment can be acquired at the cost of one good, \( R_e \) is the combined marginal product of the amount of equipment corresponding to the investment of one good in equipment.

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

\[ y = (A \Omega)^{\frac{1}{r}} \left[ \frac{\alpha}{r_s} \right]^\frac{\alpha}{1-r} \left\{ \theta_u (1 - h_s)^\sigma + (k_e^0 + \theta_s h_s^0)^\frac{\sigma}{\rho} \right\} \frac{1}{p_t}. \]

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

\[ k_e = \left[ \frac{w_s}{\theta_s r_e} \right]^{\frac{1}{\sigma}} h_s. \]

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

\[ \pi \equiv \frac{w_s}{w_u} = (k_e^0 + \theta_s h_s^0)^\frac{\sigma-\rho}{\sigma} \frac{\theta_s}{\theta_u} \frac{h_s^{\rho-1}}{(1-h_s)^{\sigma-1}} \]

\[ = \left( \left[ \frac{w_s}{\theta_s r_e} \right]^{\frac{\sigma}{\rho}} + \theta_s \right)^{\frac{\sigma-\rho}{\rho}} \frac{\theta_s}{\theta_u} \left[ \frac{h_s}{1-h_s} \right]^{\sigma-1}, \]

where the last equality is obtained by invoking expression (14). By using first-order conditions (9) and (10), we can also find an equilibrium expression for the labor share of output, \( \chi \),

\[ \chi \equiv \frac{w_u (1 - h_s) + w_s h_s}{A F(k_s, k_e, h_s)} = (1-\alpha) \frac{\theta_u (1 - h_s)^\sigma + (k_e^0 + \theta_s h_s^0)^\frac{\sigma-\rho}{\rho} \theta_s h_s^0}{\theta_u (1-h_s)^\sigma + (k_e^0 + \theta_s h_s^0)^\frac{\sigma}{\rho}} \]

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

where the last equality is obtained by twice invoking expression (15).

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

**Claim 1.** For a given labor composition \( h_s \) and rental rates \( \{r_s, r_e\} \), let the equilibrium be denoted \( \{k_s, k_e, w_s, \pi, \chi\} \). Then by varying the initial parameters \( \{\theta_u, \theta_s\} \), we can compute a continuum of equilibria, \( \{\hat{k}_s, \hat{k}_e, \hat{w}_s, \pi, \chi\} \), where the skill premium and the labor share of output 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 (13) and (14) for output and equipment, we confirm that first-order condition (10) for the employment of skilled labor continues to be satisfied at the unchanged value $h_s$ under the alternative equilibrium wage $\hat{w}_s$ and parameter values $\{\hat{\theta}_u, \hat{\theta}_s\}$. Second, given the constant value of the skilled labor share, we confirm that the skill premium in expression (15) is also the same across the equilibria. Third, given the constant values of both the skilled labor share and the skill premium, we confirm that the labor share of output in (16) is also unchanged, i.e., we verify that $\hat{\theta}_u \hat{\theta}_s^{-\sigma} = \theta_u \theta_s^{-\sigma}$.

4.2 Calibration procedure

Our objective is to calibrate a model for each country based on its labor skill composition in 1970, and then simulate the calibrated model with the time series of labor inputs over the period 1970-2002 to predict the evolution of the skill premium. Besides the actual skilled labor share and the associated skill premium measure in 1970, our calibration procedure draws upon earlier studies regarding observations on the average annual real return ($r$), depreciation rates ($\delta_u, \delta_s$), equipment price movements ($p_{t+1}/p_t$), and the labor share of output ($\chi$), and estimates of technology parameters ($\alpha, \sigma, \rho$). The data on the real return, depreciation rates and equipment price movements imply rental rates on structures and equipment, as given by equations (11) and (12). Except for the fact that these average annual values and parameters reflect economic outcomes over a long period of time, our calibration with respect to the remaining parameter values ($\Omega, \theta_u, \theta_s$) is guided by the single observation of a country’s skilled labor share in 1970 (and the associated skill premium measure).

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

1. For an arbitrary value of $\Omega$, we can compute the labor share of output as a function of the parameters $\theta_u$ and $\theta_s$. Such a mapping is illustrated in Figure 4 for one particular value of $\Omega$. Given a calibration value for the labor share of output, say $\chi = 2/3$, the bold curve in Figure 4 depicts pairs of parameter values $\{\theta_u, \theta_s\}$ that produce the targeted value $\chi$.

2. Both the labor share and the skill premium are constant along the bold curve in Figure 4, as described in Claim 1. Hence, given the calibration value of the labor share, the bold curve in Figure 4 represents a mapping from the parameter $\Omega$ (that is held fixed in the graph) and 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 5.

3. Given a calibration value for the skill premium, Figure 5 pins down our choice of parameter $\Omega$. Given that value of $\Omega$, the associated Figure 4 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, say a normalization with $w_s = 1$. 

12
Figure 4: Labor share of output, $\chi$, 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.

5 Model simulation 1970-2002

5.1 Parameterization

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

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

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

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

c) Equipment prices have fallen annually by 5 percent over the period 1970-1992, based on a time series constructed by Gordon (1990) until 1983 and by KORV thereafter. Hence, we assume an annual growth rate of equipment prices of $p_{t+1}/p_t = 0.95$.

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
Figure 5: Skill premium, $\pi$, as a function of the parameter $\Omega$, given that the labor share of output is held constant, $\chi = 2/3$.

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

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

The premises in items b)–d) yield rental rates on structures, $r_s = 0.09$, and equipment, $R_e = 0.209$, as given by equations (11) and (12). We assume that these rental rates are the same across the U.S. and Sweden.

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

As a sensitivity analysis, we recalibrate the Swedish model below using Swedish data on capital structures’ share of income and find that our results are not much affected by the alternative parameterization. Therefore, as a benchmark model, we prefer as far as possible to keep the underlying premises the same across the models for the U.S. and Sweden, to demonstrate that the observed divergence in skill premia is indeed driven by observed labor inputs in the private sector. Though, there is one dimension in which our adoption of a key estimate differs across the models:
Concerning the parameter $\sigma$, we adopt Katz and Murphy’s (1992) estimate for Sweden, $\sigma_{\text{KM}} = 0.291$ and KORV’s estimate for the U.S., $\sigma_{\text{KORV}} = 0.401$. Hence, the elasticities of substitution between unskilled and skilled labor in Sweden and the U.S. are 1.41 and 1.67, respectively.

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

Premises a)–f) combined with data on each country’s skilled labor share in 1970 (and its associated skill premium measure) allow us to select the remaining parameters $\{\Omega, \theta_u, \theta_s\}$ according to the calibration procedure in section 4.2.\textsuperscript{11}

### 5.2 Predicted skill premium

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

Figure 6 displays the actual and predicted skill premium evolution in Sweden and the U.S. The predictions are remarkably close to the actual evolution, given that the model is calibrated using data from 1970 only.

### 5.3 One different premise between Sweden and the U.S.

The skill premium evolution in Sweden and the U.S. is surprisingly well explained by a common framework that shares five of our six premises underlying the parameterization in section 5.1. The only difference, premise f), concerns the elasticity of substitution between unskilled and skilled labor. In particular, our model for Sweden adopts Katz and Murphy’s (1992) estimated elasticity of 1.41, while our model for the U.S. uses KORV’s estimated elasticity of 1.67. This difference in elasticities might seem small relative to the prevailing range of estimates in the empirical literature. For example, Katz and Autor (1999)

\textsuperscript{11}KORV do not report the values of parameters $\{\mu, \lambda, \psi_u, \psi_s\}$ in production function (5), except for mentioning that $\psi_s$ is fixed as a normalization. In our calibrations, the implied values of those parameters are $\{\mu, \lambda, \psi_u\}_{\text{U.S.}} = \{.5625, .8610, 2.3736\}$ and $\{\mu, \lambda, \psi_u\}_{\text{Sweden}} = \{.7112, .9784, 1.3097\}$, given the normalization $\psi_s = 1$. 

15
Figure 6: Skill premia in Sweden and the U.S. The solid lines refer to actual skill premia and the dashed lines depict benchmark simulations. The dotted line for Sweden is an alternative simulation with parameters \( \{ \alpha, \rho \} = \{0.186, -1.25\} \).

We report that cross-sectional and time series studies lend support to a range of 0.5-2.5. However, in the parameter space of our structural framework, the difference between an elasticity of 1.41 versus 1.67 is big.

To provide a metric for the significance of the postulated difference in the elasticity of substitution between unskilled and skilled labor in Sweden versus the U.S., we explore the implications of varying that elasticity. Specifically, keeping the same premises as in the benchmark models except for varying the parameter \( \sigma \), we calibrate a continuum of new models for Sweden and the U.S., respectively. These models are then simulated to predict the skill premium in 1990 and 2002, as depicted in Figure 7 by the dashed lines for Sweden and the solid lines for the U.S. The simulations of the benchmark models for Sweden and the U.S. are identified by circles. The simulated skill premia should be compared to the actual outcomes in 1990 and 2002, as represented by the dotted lines (and also marked with crosses at the benchmark elasticities of substitution). The distance between a circle and a cross for a given country and year is the same as the corresponding difference between the actual and simulated skill premium for that country and year in Figure 6. As we already know, the benchmark model for Sweden with an elasticity of 1.41 almost perfectly explains the Swedish skill premium in 1990 and 2002. But if we instead had assumed an elasticity of 1.67, the model for Sweden would have erroneously predicted that the Swedish skill premium in 1990 had reverted to its 1970 level, followed by a further explosive increase in 2002. Similarly, the benchmark model for the U.S. with an elasticity of 1.67 explains fairly well the actual skill premium evolution in the U.S. But if we instead had assumed an elasticity of 1.41, the model for the U.S. would have counterfactually predicted a depressed U.S. skill premium in both 1990
Arguments for why the elasticity is lower in Sweden than in the U.S.: First, Bergström and Panas (1992) provide empirical evidence of a relatively low elasticity of substitution between unskilled and skilled labor in Swedish industries. Second, the skilled fraction of the labor force is much lower in Sweden as compared to the U.S. ...

To demonstrate robustness of our analysis in another dimension of the parameter space, we consider the fact that the national income accounts suggest a higher capital structures’ share of output in Sweden, $\alpha = .186$ in the private sector (which is, incidentally, similar to the share of 0.18 in the aggregate economy). We now adopt this alternative parameter value $\alpha = 0.186$ but also adjust the elasticity of substitution between skilled labor and capital equipment, $1/(1 - \rho)$, to accommodate this different technology specification. In particular,

Figure 7: Sensitivity analysis with respect to the elasticity of substitution between unskilled and skilled labor, $1/(1 - \sigma)$. Keeping the same premises as in the benchmark models except for varying the parameter $\sigma$, we calibrate a continuum of models for Sweden and the U.S., respectively. These models are then simulated to predict the skill premium in 1990 and 2002, as depicted by the dashed lines for Sweden and the solid lines for the U.S. The simulations of the benchmark models for Sweden and the U.S. are identified by circles. (Recall that the elasticities of substitution in Sweden and the U.S. are assumed to be 1.41 and 1.67, respectively.) The actual skill premia in 1990 and 2002 are represented by the dotted lines (and also marked with crosses at the benchmark elasticities of substitution).

11Bergström and Panas (1992) estimate separate elasticities for 4 years (1963, 1968, 1974 and 1980) and for 8 manufacturing industries in Sweden. The estimates are fairly stable across time for each industry and the average elasticity in each of the 8 industries are 0.36, 0.47, 0.91, 1.25, 1.26, 1.31, 1.45, and 3.30.
we calibrate the value $\rho = -1.25$ that allows us to retain the good match with the actual Swedish skill premium in 2002. It should be noted that this value is well within the range of the empirical estimates (see Hamermesh 1993).

The dotted line in Figure 6 depicts this alternative parameterization of the model for Sweden, and we see that it explains the evolution of the Swedish skill premium as well as the benchmark model. But while this latter recalibration utilizes information about actual outcomes in both 1970 and 2002, recall that the benchmark parameterization is only based on the labor skill composition in 1970.

5.4 Other implications

The model has several other implications. The first implication concerns the labor share of output. As noted by KORV, the specification of the production function does not guarantee a constant labor share of output. But the model for the U.S. yields a labor share that remains at roughly two-thirds of output throughout the period. For Sweden on the other hand the model predicts a substantial fall in the labor share of output, whereas it remains fairly stable in the data, though somewhat lower in the 1990’s compared to the 1970’s. While a serious flaw it can easily be corrected. In our alternative parametrization, the fall in the labor share mirrors that in the data and at the same time the evolution of the skill premia predicted by the model still closely conforms with that observed in the data.

The second implication concerns productivity growth. Essentially all productivity growth can be explained by equipment-specific technological change, as earlier argued by Greenwood, Hercowitz and Krusell (1997). In the model for the U.S., output per worker grows by 2 percent per year, which is a tiny bit less than the 2.2 percent observed in U.S. data.13 For Sweden, the model predicts higher productivity growth, whereas actual growth has been somewhat smaller.14 Again this can be corrected using the alternative parametrization.

The third implication concerns the growth rate of equipment. The model for the U.S. predicts that the stock of equipment in efficiency units grow at an annual rate of 7.7 percent which is very similar to the 7.4 percent reported in KORV based on Gordon (1990) capital series. In the model for Sweden suggests that equipment grew even faster. How does this fit the facts? While there exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon’s (1990), Statistics Sweden reports data on capital stocks measured using similar methods as in the U.S. NIPA. KORV reports that U.S. equipment, as measured in official NIPA, grew at an annual rate of 3.4 percent between 1980-92. According to Statistics Sweden,15 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. Moreover, the model give predictions of equipment per worker. According to UN population data, the

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13 Real GDP grew by 3.2 percent on an annual basis (see Table 1.1.1 at www.bea.gov), and according to UN population data, the US workforce (age 15-64) grew an annual rate of 1 percent.

14 Data from Statistics Sweden shows that GDP grew by 2 percent annually and UN population data, shows that the Swedish workforce (age 15-64) grew at an annual rate of 0.3 percent.

Swedish workforce (age 15-64) grew at 0.4 percent annually during this period, while the U.S. counterpart grew at an annual rate of 1 percent. This suggests that equipment per worker may have grown faster in Sweden than in the U.S..

Table 3: Model simulations

<table>
<thead>
<tr>
<th></th>
<th>USA benchmark</th>
<th>Sweden benchmark</th>
<th>Sweden alternative</th>
</tr>
</thead>
<tbody>
<tr>
<td>Labor share of output in 2002</td>
<td>0.645</td>
<td>0.556</td>
<td>0.630</td>
</tr>
<tr>
<td>(equal to 0.667 in 1970)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Annual growth rate of output</td>
<td>2.0%</td>
<td>2.9%</td>
<td>1.6%</td>
</tr>
<tr>
<td>(and of structures)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Annual growth rate of equipment</td>
<td>7.7%</td>
<td>9.8%</td>
<td>7.7%</td>
</tr>
</tbody>
</table>

5.5 Earlier study using the framework

This framework has recently been applied on Swedish data by Lindquist (2005), though he offers a starkly different view on how the Swedish skill premium has evolved during the last three decades. When estimating the production function proposed by KORV (equation 5), Lindquist uses the university wage premium among male white-collar workers in manufacturing, mining and construction (uncorrected for age differences) as a proxy for the economy-wide skill premia. As noted above (section 3.1) this is a poor proxy and indicates too little skill premium compression; in 2002 this measure is back to its 1970 value whereas we showed, using data for all individuals, that it is some 20 percent below. It is therefore surprising to note that his estimates for $\sigma$ (around 0.3) and $\rho$ (around 0.93) are very similar to those used in this paper.

6 Discussion

6.1 Public sector employment

There is an important qualification to our measure of labor input. It is measured using private sector employment only, not aggregate employment. This is immaterial for the study of U.S. skill-premium but of crucial importance when explaining Swedish labor market outcomes. The reason for focusing on private sector employment is our belief that private consumption and government consumption are not perfect substitutes from the perspective of the households in the economy. Government expenditures must be financed by taxation, and taxes cause private valuations of taxed goods and services to differ from their true social costs. Subsidizing publicly produced consumption goods for private use may therefore lead to excessive consumption of these goods, and distort the demand for labor services used in the production.

16 This differs from, for example Edward Prescott’s analysis of why aggregate hours worked in Europe is lower than in the U.S. (see Ely lecture...)
Figure 8: Public employment share.

Figure 8 shows the share of public sector employment in Sweden and the U.S. between 1970 and 2002. In 1970, the countries were remarkably similar, with public sector employment around 22 percent of total employment. By the mid 1980’s the countries looked very different. In the U.S. the share had increased by one or two percentage points. In Sweden, on the other hand, public sector employment had increased by more than 50 percent and amounted to 35 percent of total employment. Of course, this is of no importance if the skill composition evolved in similar fashion in the private and the public sector. In the U.S., this is indeed the case. Between 1970 and 2002, the logarithmic change in ratio of skilled to unskilled was of similar magnitude in private sector employment and in total employment: 1.08 compared with 0.90. In Sweden on the other hand, the ratio of skilled to unskilled increased by much more in the private sector. The logarithmic change between 1970 and 2002 was 1.68 in private employment and 1.22 in total employment.

The mirror image is that the public sector employed an increasing share of low-skilled workers over time. What kind of jobs were these workers employed to perform? Björklund and Freeman (1997) notes that essentially all employment growth during this period was in services provided by the local government. Underlying this development was the rapid growth of publicly provided day care for pre-school children. In 1970, 7 of the 22 percent employed in the public sector worked in public day care. In 1990, half of public employment was in public day care. In 1991-92, public expenditures for families with pre-school children (parental leave, day care etc) were almost 3.5% of GDP.

In the words of Sherwin Rosen, Sweden "monetized the household sector of its economy by subsidizing publicly for privately produced household services on a grand scale... . The welfare state encourages extra production of household goods and discourages production of material goods. Too many people provide paid household (family) services for other people, and too few are employed in
the production of material goods. This is what explains the growth of local government employment."

Thus, as Lindbeck (1997) notes, "the huge increase in public sector employment kept up the demand for low and medium-skilled." As a result the skill composition in the private sector was distorted. It is important to note however, that the public sector did not hire an disproportionate number of low-skilled workers, but rather that it greatly increased its share of such workers during this period.

To investigate the importance of using private sector employment, we use total employment as a sensitivity check. Figure 9 shows that the evolution of the actual skill premia are unchanged, which is comforting, and suggesting that the assumption of free labor mobility is reasonable. 17

Figure 10 shows that the predictions based on total employment using the model with capital skill complementarity. For the U.S., the results do not change much. This reflects the fact that the skill composition in the U.S. did evolve similarly in the private and the public sectors. For Sweden, the picture is very different. The model predicts a Swedish skill premia in 2002 just short of its 1970 value, and not 20 percent below as in the data.

The results are even more extreme using the Katz-Murphy model. It counterfactually predicts that the Swedish skill premia should have risen by 20 percent.

6.2 Implications for Europe

Earlier observers have imputed importance to the expansion of the public employment in Sweden. Lindbeck (1997) as an example writes that "it is often

17 In fact, economic models with free labor mobility imply not only that changes in the skill premium should be the same across sectors, but also that wages are equalized across sectors for the same type of labor. In our study, we do not address the common assertion of seemingly lower levels of earnings in the public sector as compared to the private sector.
argued that the low unemployment rate in Sweden during the 1970’s and 1980’s was a result of the expansion of permanent public-sector employment.⁴⁹

At the same time it is fair to say that observers have in general been at a loss to rationalize the dramatic decline in the Swedish skill premium. For example, Edin and Topel (1997) suggest that the answer should be sought in a conspiracy theory involving the trade union and the employers’ association.¹⁸

Our paper demonstrates that the observed quantities of skilled and unskilled labor in the private sector of Sweden are consistent with the observed skill premium compression, even in a framework where labor is paid its marginal product.

There are two different views how this could have come about. One view is that Swedish trade union has set the skill premium and the government, acting as an employer of last resorts, has accommodated this wage structure by expanding public sector employment. Another view is that public sector expansion was an exogenous factor determined by political preferences on for example day care services. The realized skill premium is then just reflecting market conditions where unskilled workers have become relatively scarce in the private economy of Sweden.

It is important to emphasize that our analysis does not explain how the quantities of skilled and unskilled labor come about and there are two shortcomings worth mentioning. First, as in Krusell et al (2000), 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.¹⁹ Second,

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⁴⁹See also Hibbs (1990) and Davis and Henrekson (2005).

¹⁸It 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 cost and return side of investments in education have been “socialized” in form of free education with generous student benefits and a labor income tax system that has been highly progressive, if
our analysis does not contradict arguments that ascribe an important role to Swedish trade unions which might have entailed a direct influence on relative wages as long as public sector employment has accommodated any resulting market imbalances.

The conclusion of our analysis is merely that the observed skill premium and the skill composition of the Swedish labor force are fully consistent with labor being paid its marginal product.

We believe that this study of the compressed skill premium in Sweden might further our understanding of the European employment situation. Most European welfare programs, such as long-term unemployment benefits, early retirement, and disability, have some form of ceiling on benefit levels. This implies that the low-skilled are more likely to face higher replacement rates, and this therefore results in a disproportional withdrawal of unskilled labor from European employment. If economic growth in the last couple of decades have been driven by equipment-specific technological change coupled with skill complementarity, and indeed studies by Grenwood et al (1997) and Sakellarisa and Wilson (2004) suggests that this is the case, it might very well provide an important explanation to the worsening employment situation in Europe. It is the combination of such a technological development and the existence of generous social programs that has then enabled the furlough of unskilled labor in Europe which has kept the skill premium from rising. This view provides a structural explanation to the European unemployment dilemma described by Ljungqvist and Sargent (1998) who ascribe the outbreak of high European unemployment in the 1980s to the outbreak of economic turbulence.

References


not outright confiscatory, during long periods of time.

20 As described by Ljungqvist and Sargent (1998), their framework is meant to capture skill accumulation in form of work experience among blue-collar workers. To introduce our concept of a skill premium, their model would have to be augmented to include college-educated white-collar workers. In such a framework with caps on absolute benefit payments, we conjecture that low-skilled blue-collar workers with their high replacement rates would suffer disproportionately from long-term unemployment as modelled by Ljungqvist and Sargent (1998). Hence, a quantitative study of such a model might offer explanations to both the higher incidence of long-term unemployment and the compressed skill premia in Europe as compared to the United States.


Appendix A  Data Construction

This appendix describes how we construct our measures for skill premium and labor input. 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 covers 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 premium and labor input using two different samples. The supply sample include 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 others), 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 services, education, postal services and government administration). While health services and to some extent educational services 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.

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 for 1970-2002 as weights.

All variables used are reported in Tables A1-A3.


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 50916 in 1970, 74260 in 1980, 68491 in 1990 and 60332 in 2000.

In the wage sample we follow Autor et al (2005) and exclude (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 has been made, see below) is less than $67 in 1982 dollars (we have used CPI when transforming earnings into 1982 dollars), or whose hourly wage exceed $35th the topcoded value of weekly earnings. The number of remaining observations are 29009 in 1970, 43260 in 1980, 49260 in 1990 and 45332 in 2000.

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

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

Prior to March 1989, wage and salary income is reported in the CPS as a single variable, that is topcoded at 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. Beginning in 1996, topcoded earners are assigned the mean of all topcoded earners (averages when individuals are grouped according to sex, race and worker status). Figure 11 shows that the share of individuals that are topcoded has increased over time (and that each

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21 We have investigated the following alternative specifications with unchanged results; (i) individuals whose weekly earnings are less than 40 times half the hourly minimum wage are included but their wages are imputed to equal half the minimum wage, (ii) individuals whose weekly earnings are less than 40 times the hourly minimum wage are excluded.
We have therefore investigated four different topcoding procedures. In all of them, we topcode primary and secondary earnings separately before summing them and we recensor primary and secondary earnings for the years 1996-2002 to their top nonrecoded values. The first procedure involves no further topcoding. In the second procedure, we follow Autor et al. (2005) and multiply topcoded values by 1.5. In the third procedure, we follow Card and DiNardo (2002), and recensor primary and secondary earnings after 1989 to their 1988 topcoded values of $99,999 and $25,000 respectively. In the fourth procedure, we assume that the upper tail of the wage and salary distribution is Pareto distributed. For each year we multiply the topcoded value with the ratio of the total income of topcoded individuals implied by the Pareto distribution and the observed total income. In particular, we proceed as follows. First, let the year-specific cumulative distribution function for the Pareto distribution be given by

\[ F(x_t) = 1 - \frac{b_t^{a_t}}{x_t^{a_t}} \]

where \( x \) denote wage and salary, and \( a_t \) and \( b_t \) are parameters. Then let \( x_{c,t} \) denote the topcoded value in year \( t \). Let \( x_{90,t} \) denote the wage and salary that solves

\[ F(x_{90,t}) = 0.9 \ . \]

For each year we can the calibrate the distribution by finding the \( a_t \) and \( b_t \) that

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\(^{22}\) The top nonrecoded values for primary and secondary earnings are $150,000 and $25,000 respectively before 2003 and $200,000 and $35,000 thereafter.

\(^{23}\) Recent papers that have used the Pareto interpolation technique include Frenberg and Poterba 1993, Piketty 2003 and Piketty and Saez 2003.
solves
\[ 1 - \frac{b_{a,t}^{c,t}}{x_{a,t}^{90,t}} = 0.9 \]
and
\[ \frac{b_{a,t}^{c,t}}{x_{c,t}^{a,t}} = y_t \]
where \( y_t \) is the fraction of individuals that are topcoded in the data.

Then let \( X_{c,t} \) denote the sum of the topcoded individuals’ income in the data (i.e., when topcoded). That is
\[ X_{c,t} = (1 - F(x_{c,t}))x_{c,t} = \frac{b_{a,t}^{c,t}}{x_{c,t}^{a,t}}x_{c,t} \]
Let \( X_t \) denote the sum of the topcoded individuals income implied by the Pareto distribution
\[ X_t = \int_{x_{c,t}}^{\infty} \frac{b_{a,t}^{c,t}}{x_{a,t}^{a,t}} x_t dx = \frac{a_t}{(a_t - 1)} \frac{b_{a,t}^{c,t}}{x_{c,t}^{a,t}}x_{c,t} \]
Thus the ratio of the implied and the observed income is given by
\[ \frac{X_t}{X_{c,t}} = \frac{a_t}{(a_t - 1)} \]
For each year we then multiply the topcoded value with \( \frac{a_t}{(a_t - 1)} \). From 1989 and onwards we use primary earnings to derive \( \frac{a_t}{(a_t - 1)} \) and then use this value for both primary and secondary earnings. Figure 12 shows how this factor is approximately 1.5 until 1985 and then gradually increases towards 1.8 in the end of the sample.

Figure 13 shows the U.S. skill premium for the four different procedures for topcoding. The overall picture is rather similar. However, the extent to which the 1970’s was characterized by a decreasing skill premium depends on how individual wages are topcoded. Also, procedures 1, 2, and 4 predict a similar increase in the skill premium during the 1980’s and 1990’s of approximately twenty percent. The third procedure on the other hand predicts that the increase in the skill premia was an episodic event of the 1980’s and that it remained fairly stable during the 1990’s. This is due to the recensoring of earnings to their 1988 values. If skilled individuals are overrepresented among those that are recensored, then this procedures automatically biases the skill premia downwards. And indeed they are. In 1989, 0.56 percent of the skilled and 0.22 percent of the unskilled had their primary earnings censored. In 2002, the share of censored skilled and unskilled individuals were 3.42 percent and 0.92 percent respectively.\footnote{For secondary earnings, the shares for skilled and unskilled respectively were 0.17 percent and 0.09 percent in 1989 and 0.51 percent and 0.42 percent in 2002.}


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, (iii) individuals
Figure 12: Topcode-factor when assuming a Pareto distribution.

Figure 13: Skill premium in United States for various topcoding procedures.
without education or industry classification. The number of remaining observations are approximately 3.5 million in 1970, and 3.9 million in the period 1990-2002.

We construct the wage sample by excluding (i) the self-employed, and (ii) individuals who work less than 20 hours per week. Hours per weeks are reported in the following intervals: 1-19, 20-35, >35 in 1970, and 1-15, 16-19, 20-35, and >35 in 1990. The main reason for including individuals in the interval 20-35 in the wage sample is that average hours worked per week by women is around 30 hours throughout the period (see below). The number of remaining observations are approximately 3.3 million in 1970, and 3.7 million in the period 1990-2002.

The classification scheme for education is 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 classify both groups as high school graduates. Second, obtaining a traditional Swedish college degree requires 3-5 years of studies depending on field of specialization. We thus let college graduate include individuals who have completed a 3 year university or college degree.

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, and SN192 codes (0-64110, 64204-73000, 74111-75000, and 92000-95001) in the LOUISE database for 1990-2002.

A.3 Group-specific wages and hours worked

Let \( d_{i,t} = 1 \) if individual \( i \) in year \( t \) belongs to the wage sample. For each individual \( i \), we use observations on the number of hours worked per week, \( h_{it} \), the number of weeks worked per year, \( wk_{it} \), annual earnings, \( y_{it} \), and the sampling weight \( \mu_{it} \). For each group \( g \in G \) we obtain measures for the hourly wage rate \( w_{gt} \), as

\[
w_{gt} = \frac{\sum_{i \in g} y_{it} \mu_{it} d_{i,t}}{\sum_{i \in g} wk_{it} \mu_{it} d_{i,t}}.\]

For each group we then calculate measures for average annual hours worked \( l_{gt} \) as

\[
l_{gt} = \frac{\sum_{i \in g} wk_{it} h_{it} \mu_{it}}{\mu_{gt}},\]

where \( \mu_{gt} = \sum_{i \in g} \mu_{it} \).

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 (\( \mu_{it} = 1 \)). Second, in the U.S. data beginning in 1976 hours worked refer 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'. We assume that individuals are perfect substitutes within a group and calculate the supply of hours for these individuals as

\[
h_{it} = \frac{y_{it}}{w_{gt} wk_{it}}, \quad i \in g \text{ and } d_{i,t} = 0.
\]
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 we have observations on hours worked per week, \( h_{gt} \) which we multiply by 52 to obtain annual hours worked. Moreover, hours worked is reported in intervals. In 1970 the intervals are 1-19, 20-34 and 35+ hours per week. In 1990 the intervals are 1-15, 16-19, 20-34, 35+. We have assigned hours per week using midpoints, assuming 43 hours per week for 1970 and 40 hours per week for 1990 in the upper interval.\(^{25}\) A potential problem with this procedure is that the implied average number of hours per week do not exactly match the ones reported by Statistics Sweden (see www.scb.se and Pohjolassa 1983). In the data the average hours worked in 1970 was 39.3 hours per week, with men working 43.5 and women 32.8 hours per week. In 1990 the equivalent numbers were 37.4, 40.4, and 33.9, respectively. Using midpoints implies average working hours of 39.9, 42.6, and 35.8 for 1970 and 36.4, 38.7, and 33.9 for 1990. That is, midpoints overestimates the working hours of women in 1970 and underestimates the working hours of men in both 1970 and 1990.

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. We use the following procedure to assign individual hours in 1991-2002. For each group \( g \), we use 1990 data to obtain the group-specific distribution of hours worked across intervals. We assume that these distributions remain constant after 1990. For each year we sort individuals according to their annual earnings. Individual hours is then then assigned by imposing the 1990 distribution in descending order.

### A.4 Skill premium and efficiency unit of labor input

The groups are then aggregated into two classes; \( j = s \) (skilled) and \( u \) (unskilled). For each class, total labor supply is given by

\[
L_{jt} = \sum_{g \in G_j} l_{gt} \bar{w}_g h_{gt}
\]

where \( \bar{w}_g \) are average group wages for 1970-1990.\(^{26}\) Moreover, the average hourly wage for each class is given by

\[
W_{jt} = \frac{\sum_{g \in G_j} l_{gt} w_{gt} h_{gt}}{L_{jt}}.
\]

---

\(^{25}\)We choose these 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.

\(^{26}\)We have also used 1970 group wages, \( w_{g,70} \), and 2002 group wages, \( w_{g,02} \), respectively as weights, with similar results.
A.5 Wage premium for specific professions

The identification scheme used when deriving the wage premium between specific professions is as follows. First, since part of the differences in wage premium between countries may be driven by differences in age and gender composition for the chosen professions, we control for these factors, by restricting the sample to white males in the age group 40-59. Second, in the U.S. data we identify engineers as those with CPS occupation codes 6-23 in 1970, or 44-59 in 1990. Blue-collar workers are identified as craftsmen, operators and laborers (CPS codes 400-799 in 1970, or 500-899 in 1990). Third, the Swedish observation for 1970 is based on Ståhl 81974). He reports average yearly income of 85 000 SEK for 45-55 year old engineers and of 50 000 SEK for 50 year old blue-collar workers. This suggest a wage-premium between these two groups of 1.7. Finally, the Swedish observation for 1990 is calculated using the LOUISE database. We identify engineers and blue-collar workers using ‘Nordisk Yrkesklassificering FoB-85’. Blue-collar workers are those with codes 50000-89999 and engineers are those with codes (120, 125, 130, 140, 160, 180, 220, 225, 230, 240, 250, 260, 290, 320, 325, 330, 340, 350, 360, 370, 390, 420, 425, 430, 440, 460, 490, 520, 525, 530, 540, 550, 560, 590, 620, 625, 630, 640, 650, 660, 690, 720, 725, 730, 740, 750, 760, 780, 790, 820, 825, or 850).

A.6 Tables

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<td>Sex</td>
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Table A2: Census variables used for Swedish labor input

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<td>Sex</td>
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<td>Wage and salary income</td>
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Table A3: LOUISE variables used for Swedish labor input

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<tr>
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<tr>
<td>Industry</td>
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<tr>
<td>Hours worked per week</td>
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<tr>
<td>Wage and salary income</td>
<td>ArbInk $y_{it}$</td>
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