

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 30 percent between 1970 and 1990 while the U.S. skill premium, after an initial decline in the 1970s, rose by around 8 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. In our analysis, the dramatic decline of the skill premium in Sweden is the result of an expanding public sector that has disproportionately hired unskilled labor.

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 from numerous seminar participants. Financial support from The Jan Wallander and Tom Hedelius Foundation is gratefully acknowledged.

1 Introduction

A longstanding puzzle is the diverse wage structure experiences of Western market economies in the last fifty years. A common view is that the dramatic compression of wages in some of the so-called welfare states cannot be rationalized in terms of market outcomes but is rather due to institutional factors such as the power of trade unions. We revisit the evolution of the skill premium in perhaps the starkest example of Sweden. Whereas other studies have failed to explain Swedish outcomes in terms of marginal products, we succeed when replacing commonly used measures of total labor supplies by private sector employment. Thus, we quantitatively confirm an earlier idea on how the public sector might affect relative wages of skilled and unskilled labor, with possible ramifications for other European welfare states.

As a reference, skill premia in Sweden and the U.S. were similar in 1968 when the typical college graduate earned 50 percent more than an otherwise comparable high school graduate in both countries (Edin and Topel, 1997, p. 173).¹ But outcomes diverged sharply in the next couple of decades when Swedish census data and tax records reveal an astonishing wage compression. The Swedish skill premium fell by 30 percent between 1970 and 1990, while the U.S. skill premium, after an initial decline in the 1970s, rose by around 8 percent. Since then both skill premia have increased by about 10 percentage points in 2002.

The dramatic decline in the Swedish skill premium have led many observers to question if Swedish workers are any longer paid their marginal products. For example, Edin and Topel (1997) theorize that Swedish wages have become unhinged from marginal products because of collusion between the trade union and the employers' association.² We will instead explore if marginal products of aggregates of skilled and unskilled labor can explain the evolution of the Swedish skill premium. We use a production function with a capital-skill complementarity as formulated by Krusell, Ohanian, Ríos-Rull and Violante (2000), hereafter denoted KORV. KORV showed that the historical increase in the relative supply of skilled labor in the U.S. was more than offset by skill-biased technological change, as measured by falling quality-adjusted equipment prices, which enabled them to successfully account for the rising U.S. skill premium between 1963 and 1992.

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

²In Edin and Topel's (1997, pp. 192–197) model, the egalitarian objectives of the trade union coalesce with the interests of the employers' association because the bargained “agreements did not just raise the compensation of low-wage workers; they also reduced the absolute wage of skilled workers ... delivering ‘cheap’ skilled labor to large employers.”

As in earlier cross-country studies of skill-biased change, e.g. Gottschalk and Joyce (1998), and Acemoglu (2003), we find that total employment of skilled and unskilled labor cannot explain the Swedish skill premium. But employment in the private sector can. Our analysis suggests that the dramatic decline of the skill premium in Sweden is the result of an expanding public sector that has disproportionately hired unskilled labor. The excessive hiring of unskilled labor in the public sector propagated a relative scarcity of unskilled labor in the production function of the private sector, which pushed the skill premium down – a development that eventually came to an halt in the 1990s and was partly reversed.

To illustrate the determinants of the skill premium, consider a Cobb-Douglas technology in the private sector with unskilled labor, h_u^P , and skilled labor, h_s^P , as the two inputs; that is, private-sector output is given by a production function $f(h_u^P, h_s^P) \equiv (h_u^P)^\nu (h_s^P)^{1-\nu}$ where $\nu \in (0, 1)$. Public-sector output is given by a production function $g(h_u^T - h_u^P, h_s^T - h_s^P)$, where h_j^T is the economy’s total supply of unskilled (skilled) labor for $j = u$ ($j = s$). Profit maximization in the private sector implies that the skill premium, w_s/w_u , is equal to the ratio of marginal products in the private sector,

$$\frac{w_s}{w_u} = \frac{1 - \nu}{\nu} \left[\frac{h_s^P}{h_u^P} \right]^{-1} = \frac{1 - \nu}{\nu} \left[\frac{h_s^T}{h_u^T} \right]^{-1} \frac{h_s^T/h_u^T}{h_s^P/h_u^P}. \quad (1)$$

Hence, the skill premium is inversely related to the skill ratio in the private sector, h_s^P/h_u^P . How the skill premium varies with the skill ratio in total employment, h_s^T/h_u^T , depends on the relative evolution of the skill ratio in private versus total employment. Free labor mobility causes wages to be equal across the private and public sector, but there are no market forces that guarantee equality between wages and marginal products in the public sector.³

Our quantitative analysis pursues the assertion of Lindbeck (1997, pp. 1311) that “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.⁴ As noted by Freeman, Topel and Swedenborg (1997, p. 13), essentially all Swedish employment growth in the 1970s and 1980s was in services provided by the local government,

³In Rosen’s (1997) analysis of the Swedish welfare state, there are also two production functions for private-sector and public-sector output, respectively, where the latter output is combined with households’ own time to produce and consume ‘household goods.’ Our framework enriches Rosen’s production function for private-sector output by introducing two types of labor as well as capital structures and equipment, which enables us to conduct a quantitative analysis and to study the skill premium.

⁴Besides their theory described in footnote 2, Edin and Topel (1997, p. 190) 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.”

including 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.” But with a major economic crisis in the early 1990s, both the expansion of and the past bias in favor of unskilled labor in public-sector employment came to an halt. For example, the largest and also the least skilled worker category in health care and elderly care shrunk by more than 10 percent between 1992 and 1995 and continued to decline, while the numbers of skilled staff such as medical doctors grew steadily.⁵

The reversal in the evolution of the Swedish skill premium is key for identifying and testing our CES production function in skilled and unskilled labor, capital equipment and structures. While we retain several premises from KORV’s analysis of the U.S., e.g. the average real interest rate and the rate of decline in equipment prices, there are well-known differences in education systems including a large component of high-school vocational training in Sweden (and in the rest of Europe) which is nonexistent in the U.S. where such training is instead provided at the college level.⁶ To allow for differences in education systems, we calibrate Swedish labor substitution elasticities to outcomes either before or after 1990, and then explore the model’s predictive power for the whole period 1970–2002. We find that a calibration to any subset of the data for private employment can successfully explain the skill premium over the entire period, while no amount of data can overturn the failure of total employment.

Section 2 describes the data and our compilation of aggregate skill premia for Sweden and the U.S. The skill premium indices are largely robust to decompositions in private versus total employment, and with respect to gender. We can rule out that the dramatic decline in the Swedish skill premium is due to lower ability people acquiring higher education when the fraction of skilled labor expanded greatly, because a similar large decline is experienced by workers who were employed in both 1970 and 1990 with unchanged education.

Section 3 presents the model and parameterization, followed by simulations in section 4. We show that KORV’s success in explaining the U.S. skill premium extends to yet another decade until 2002. And in contrast to Sweden, it does not matter whether total or private employment is used because the skill ratio in total and private employment evolved similarly in the U.S. Using private employment, we identify the Swedish elasticity of substitution between skilled and unskilled labor (between skilled labor and equipment) to lie in a range

⁵For our data as well as corroborating statistics from the employers’ association, see Appendix A.2.

⁶For descriptions and statistics on vocational training in high school, see Merdich, Kagehiro and Houser (1994, table 3.2), and OECD (1998, table C2.1).

centered around 1.19 (0.31) as compared to KORV’s estimate for the U.S. of 1.67 (0.67).⁷ Given those elasticities, our models for Sweden and the U.S. do not only rationalize the evolution of skill premia but also the observed constancy of labor shares over time, rates of investments and increases in labor productivity.

Our historical account of Swedish labor market outcomes in section 5 leads up to a counterfactual question: what would have happened if the Swedish public sector had not disproportionately hired unskilled labor? The model suggests that, if the relative growth rate of the ratio of skilled/unskilled labor in Swedish private employment as compared to this ratio for actual total employment had been the same as that for the U.S., the Swedish skill premium would have increased somewhat between 1970 and 1990, and would have been almost identical to that of the U.S. in 2002. Hence, seen through the lens of our model, public sector employment in Sweden has been critical for suppressing the Swedish skill premium as compared to the U.S. skill premium. We offer some concluding thoughts in section 6.

2 Data

Using the aggregation methods of Katz and Murphy (1992) and KORV, we construct skill premia and quantities for aggregates of skilled and unskilled labor. Skilled workers 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.). We compute measures for both total and private employment.⁸ Additional details are given in 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

⁷Our finding of a lower elasticity of substitution between skilled and unskilled labor in Sweden as compared to that in the U.S., is supported by the microeconomic study of Bergström and Panas (1992) who estimate relatively lower elasticities in most industries of their Swedish data set. Krueger and Kumar (2004) present a theory that attributes U.S.-Europe differences in growth experience to European education policies that favored specialized, vocational education as compared to more general education in the U.S. The thrust of their argument appears also to be consistent with our finding of a lower elasticity of substitution in Sweden.

⁸We 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.

we construct two different samples: (i) a supply sample including all workers between 16 and 70 years of age and (ii) a wage sample which is restricted to full-time workers. In each sample, we sort workers into groups according to age, sex, race, education, and private/public sector employment. We treat individuals within groups as perfect substitutes. Using the wage sample we calculate with-in group average hourly wages as the ratio between total income and total annual hours for each group. Using the supply data we calculate total annual hours for each group.

The groups are then sorted into two classes; skilled and unskilled labor, where by skilled we mean college graduates. We obtain class-specific measures by aggregating across groups. Total labor supply for each class is given by weighting hours in each group within the class by 1990 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 annual observations from government registers in 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.⁹ 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.¹⁰

⁹See Appendix A.4 for a discussion of topcoding in the U.S. data.

¹⁰The only shortcoming of the Swedish data is that while it contains individual data on hours worked for 1970 and 1990, we have to extrapolate that information for the years 1991-2002. Specifically, we assume that average hours of work in each group of workers has remained constant at its 1990 level. This is consistent with Statistics Sweden's estimates of almost constant hours worked by age and gender among the employed between 1990 and 2000 and hence, our extrapolations are not likely to undermine the advantage of otherwise having full information about the entire population of Sweden. See Appendix A.3 for more details.

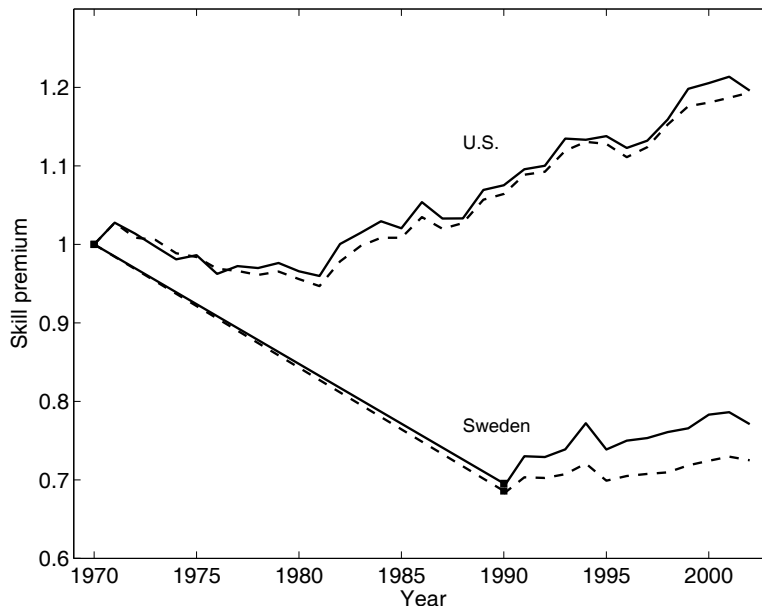


Figure 1: Skill premium in Sweden and the U.S. (normalized to one in 1970). The solid lines refer to private employment, and the dashed lines to total employment.

2.3 Trends in skill premia and labor supply

We lay out the basic facts in figures 1 through 3.¹¹ Figure 1 shows how the Swedish and U.S. skill premia evolved between 1970 and 2002. The U.S. development is well known. The skill premium deteriorated during the 1970s, but has grown rapidly since around 1980.¹² The development of the Swedish skill premia is dramatically different. Between 1970 and 1990 the skill premium fell by an astonishing 30 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 focus on private employment. Figure 3 depicts the ratio of skilled/unskilled

¹¹The first two points of Swedish time series refer to census data for 1970 and 1990, as marked by bullets and connected by straight lines, while the remainders of time series draw on the LOUISE database. 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 LOUISE as compared to the census data.

¹²Our time series for the skill premium in the U.S. is very similar to that documented by Autor et al. (2008). Recently, some authors have argued that the rise in skill premia during the 1980s was an episodic event that did not continue in the 1990s. However, as shown in Appendix A.4, this argument relies to a large extent on how the data is topcoded.

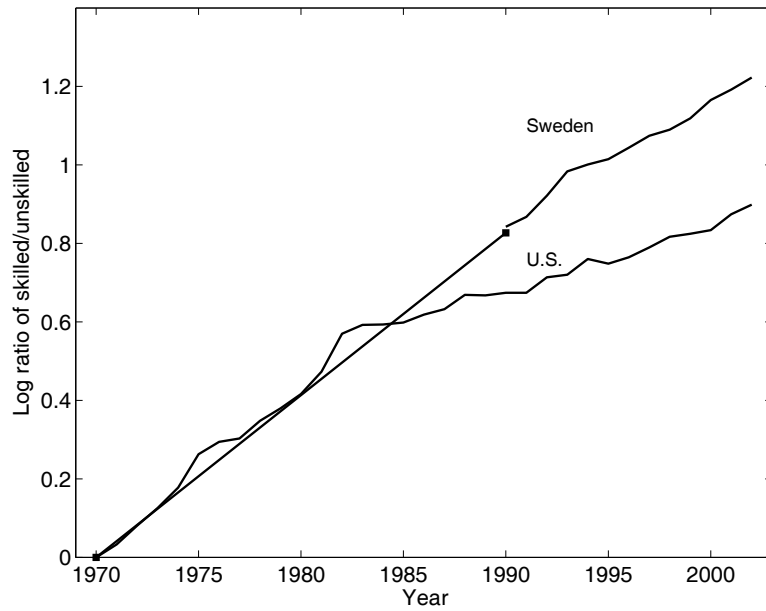


Figure 2: Ratio of skilled to unskilled labor in total employment in Sweden and the U.S. (log change since 1970).

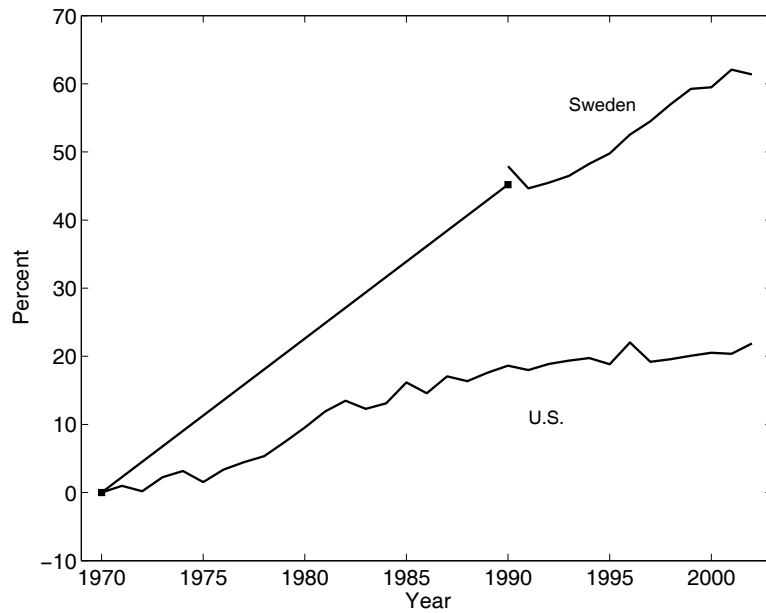


Figure 3: Ratio of skilled/unskilled labor in private employment relative to that ratio in total employment in Sweden and the U.S. (percentage change since 1970).

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 threefold of that. Consequently, the relative supply of college educated workers in Swedish private employment has grown much more than that in total employment in figure 2.

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.

2.4 Remarks about the Swedish skill premium

Appendix B addresses two potential concerns regarding the Swedish skill premium in figure 1. The first concern is that the evidence of a dramatic decline between 1970 and 1990 rests on only two years of data. Though, while only the censuses of 1970 and 1990 contain the information on education needed to compute the skill premium, we present another measure for the intermittent censuses of 1975 and 1985, based on a classification of occupations as being predominantly unskilled or skilled. Calculations of the relative wage for skilled versus unskilled occupations confirm that a dramatic decline in the skill premium has occurred, and in a rather continuous fashion. The second concern is that the computed decline in the skill premium might as well be due to lower ability people acquiring higher education when the fraction of skilled labor expanded greatly, rather than reflecting a true drop in the relative wage of skilled versus unskilled labor. We lay this concern to rest by showing that workers who were employed in both 1970 and 1990 with unchanged education, experienced a similar large decline in the skill premium. In a similar vein, since Sweden has experienced a large increase in female labor force participation, we compute the skill premium evolution for men and women separately but find no major differences.

The Swedish skill premium in figure 1 stands in stark contrast to a proxy that was used in two earlier studies that sought to explain the skill premium in terms of aggregate quantities of skilled and unskilled labor. Edin and Holmlund (1995) and Lindquist (2005) use the university-high school wage differential according to Statistics Sweden's (SCB) monthly salary data for male white-collar workers in mining, manufacturing and construction. For example, while that proxy suggests that the aggregate Swedish skill premium in 1999 was back to its 1970 value, the actual skill premium in 1999, based on census data and tax records, was not anywhere near its historical level of inequality in 1970 but, rather, more

than 20% depressed relative to 1970.¹³

3 Model

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

$$y = A k_s^\alpha \left\{ \mu (\psi_u h_u)^\sigma + (1 - \mu) [\lambda k_e^\rho + (1 - \lambda) (\psi_s h_s)^\rho]^\frac{\sigma}{\rho} \right\}^\frac{1-\alpha}{\sigma}, \quad (2)$$

where y is aggregate output, k_s and k_e are the stocks of capital structures and capital equipment, and h_u and h_s are hours of unskilled and skilled labor inputs. The only potentially time varying production parameter is A , i.e., a multiplicative neutral technology factor. The remaining time invariant production parameters are as follows. Capital structures share of income is given by α , and σ and ρ determine elasticities of substitution. The elasticity of substitution between skilled labor and capital equipment is $1/(1 - \rho)$, and the elasticity between unskilled labor and equipment or skilled labor is $1/(1 - \sigma)$.¹⁴ Income shares of different production factors are governed by μ and λ , and ψ_u and ψ_s convert hours of unskilled and skilled labor, respectively, into efficiency units.

Rates of depreciation on structures and equipment are given by δ_s and δ_e , respectively. And as in the standard growth model, KORV assume a linear transformation technology for converting output into new capital but, while the transformation rate is constant over time for capital structures, they introduce equipment-specific technological progress. In the current period, one unit of output can be converted into $1/p_e$ units of capital equipment, which is assumed to increase over time. In a competitive equilibrium, p_e becomes the relative price of capital equipment. Technological progress implies that next period's price p'_e is lower than current period's price p_e .

¹³Edin and Holmlund (1995, figure 9.4) present both the series from SCB and another estimate of the university-high school wage differential based on survey data (LNU, 'Level of Living Surveys'). The latter estimate suggests an evolution that is much more aligned with that of the aggregate skill premium in our figure 1. The LNU data are the most commonly used in studies of the Swedish wage structure. For a recent study that uses the years 1968, 1981 and 2000, see Albrecht et al. (2011, p. 1056) whose findings on wage distributions parallel our findings on the aggregate skill premium; "The first period was characterized by a dramatic equalization of wages and the second by a weak increase in inequality."

¹⁴To 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.

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 (2) as follows,

$$y = A \Omega k_s^\alpha \left\{ \theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}} \right\}^{\frac{1-\alpha}{\sigma}} \equiv A F(k_s, k_e, h_u, h_s), \quad (3)$$

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

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

$$r_s = A \alpha F(k_s, k_e, h_u, h_s) k_s^{-1}, \quad (4)$$

$$r_e = A \Gamma(k_s, k_e, h_u, h_s) (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}} k_e^{\rho-1}, \quad (5)$$

$$w_u = A \Gamma(k_s, k_e, h_u, h_s) \theta_u h_u^{\sigma-1}, \quad (6)$$

$$w_s = A \Gamma(k_s, k_e, h_u, h_s) (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}} \theta_s h_s^{\rho-1}, \quad (7)$$

respectively, where

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

The rental rate r_s on structures and r_e on equipment are such that all capital investments yield the same market-determined rate of return, say a gross rate of return equal to $1 + r$, net of physical depreciation and economic obsolescence.¹⁵ Hence, the rental rates satisfy

$$r_s = r + \delta_s, \quad (8)$$

$$\frac{r_e}{p_e} = 1 + r - \frac{p'_e}{p_e} (1 - \delta_e) \equiv \tilde{r}_e, \quad (9)$$

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 (6) and (7), we get an equilibrium expression for the skill premium,

$$\frac{w_s}{w_u} = (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}} \frac{\theta_s h_s^{\rho-1}}{\theta_u h_u^{\sigma-1}}. \quad (10)$$

¹⁵We follow KORV and ignore risk premia in our analysis of capital investments. Furthermore, to simplify the notation, our expressions are written under perfect foresight.

Appendix C.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 (4) and (5). (The implied investment rates provide an additional dimension for assessing the performance of our calibrated models, as examined in section 4.4.) Next, given the observed times series of the labor skill composition and the implied capital stocks, we compute predicted skill premia from equation (10).

The simulations are carried out in models that are calibrated to a base year (1970 in the benchmark calibrations). Besides outcomes in that base year, our calibration procedure draws upon observations in the literature on the average annual real return (r), depreciation rates (δ_s, δ_e), and equipment price movements (p'_e/p_e), which are assumed to be the same across countries. We use country-specific historical averages of the labor income share and the capital structures share of income (α). To complete the parameterization of the U.S. model, we use KORV’s estimates for the elasticity parameters σ and ρ . In the absence of comparative Swedish estimates, we need to deduce those from our data and model, as explained in section 3.3.

Regarding the remaining scale parameters, $\{\Omega, \theta_u, \theta_s\}$, Appendix C.2 describes how those are calibrated based on outcomes in the base year.

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 calibration of the models for Sweden and the U.S. share the following identical premises.

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

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

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

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

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

The premises in items a)–c) yield rental rates on structures, $r_s = 0.09$, and equipment, $\tilde{r}_e = 0.209$, as given by equations (8) and (9). 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.

3.3 Country-specific parameter values

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

To complete the parameterization of the Swedish model, we use historical averages to estimate the capital structures' share of income, $\alpha = 0.17$ in both the private sector and in the total economy; and a labor share of income equal to 0.61 in the private sector and

0.68 in the total economy. Sensitivity analyses inform us that our conclusions are robust to the choice of these factor income shares.¹⁶ Regarding elasticity parameters σ and ρ , we lack good estimates for Sweden and we choose instead to deduce these from our data and model. Given that the Swedish skill premium declined between 1970 and 1990, and then partially recovered over the next decade, it is natural to split the Swedish data in two subperiods for calibration and simulation purposes. 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 1990) 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. To anticipate our results, we find that private employment performs well in all exercises while total employment fails to predict the evolution of the Swedish skill premium.

The simulations for Sweden are performed as follows. For each calibration period and for each measure of employment (private or total), we conduct a sensitivity analysis by considering different pairs (ρ, σ) . In particular, for each ρ that belongs to a set that we have picked, we compute the value of σ that best captures the observed skill premium in the calibration period (using the least-squares metric). Given each such pair (ρ, σ) , we then simulate the model over the whole period 1970–2002. Our set of ρ is picked so that the corresponding elasticities of substitution between skilled labor and equipment span the interval $[0.1, 0.8]$. As we will show, this interval is enough to give both private and total employment the best chances of predicting the observed skill premium. Note also that the interval falls within the range of empirical estimates, as reviewed by Hamermesh (1993).

4 Simulations 1970–2002

In this section, we confirm that the explanatory power of KORV’s earlier analysis of the U.S. skill premium extends to yet another decade until 2002. When applying KORV’s model

¹⁶For details on the computations of factor income shares in Sweden, see Appendix A.5. As a sensitivity analysis we have set $\alpha = 0.1$ or $\alpha = 0.2$, and our results were largely unchanged. We have also set the labor income share to $2/3$ as for the U.S. economy. Again, our findings were robust. These and the other sensitivity analyses in the paper are available from the authors upon request.

to Sweden, we find that private employment, but not total employment, can explain the evolution of the Swedish skill premium. We end by looking at additional implications of the simulations, all of which lend support to the theoretical framework.

4.1 U.S. skill premium

We use the calibrated model and the observed time series for the labor skill composition over the period 1970-2002 to predict the U.S. skill premium. Figure 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 to calibrate the model: the base year 1970. We conclude that the explanatory power of KORV's earlier analysis of the U.S. extends to our more recent data set and, it makes not much of a difference whether we use private or total employment. The latter finding we might have anticipated from figure 3 in which the ratio of skilled to unskilled labor has developed similarly in private and total employment, respectively.

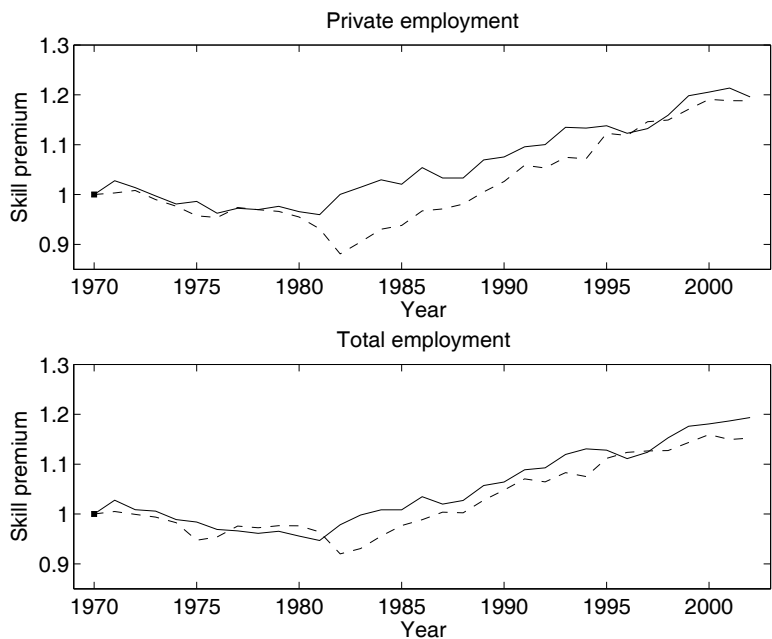


Figure 4: U.S. skill premium in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

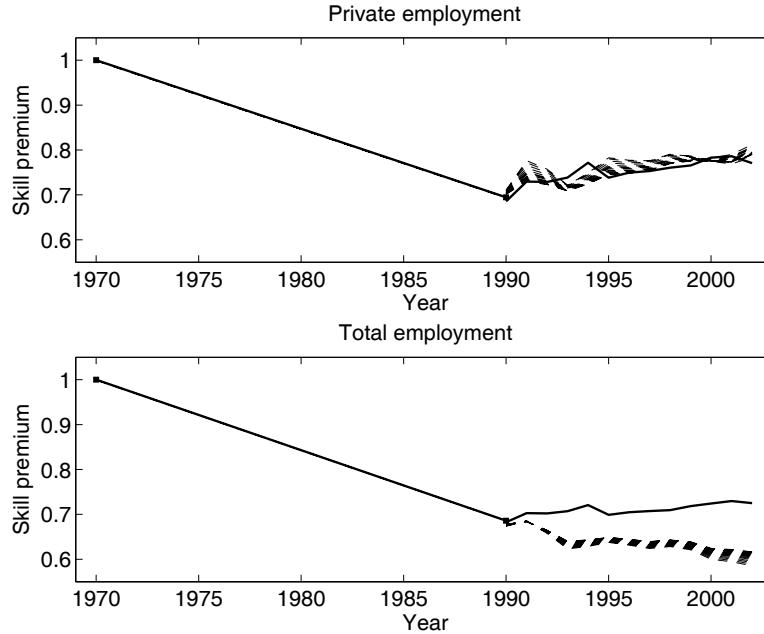


Figure 5: Swedish skill premium, given calibration period 1970-1990, in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

4.2 Swedish skill premium

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., 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’ period 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 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 period 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

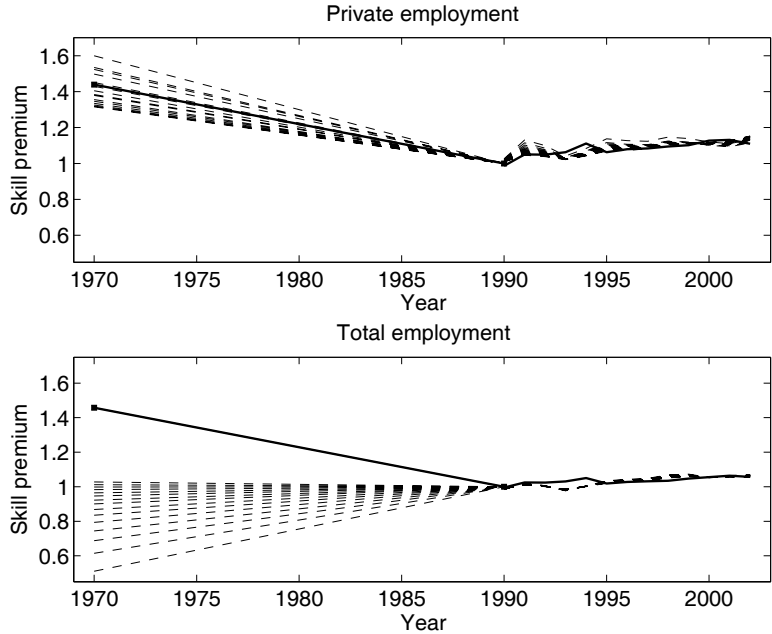


Figure 6: Swedish skill premium, given calibration period 1990-2002, in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

skill premium. The performance of the latter model based on total employment cannot be improved upon by considering elasticities outside of the stipulated interval of elasticities.¹⁷

In figure 7, we use the whole period 1970–2002 to calibrate the substitution elasticities. As compared to figures 5 and 6, the upper panel based on private employment in figure 7 does not look much different. That is, 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 procedures. As shown in the lower panels of figures 5 and 6, total employment cannot

¹⁷Since the upper simulated path in the lower panel of figure 6 predicts a skill premium in 1970 that is almost equal to its 2002 value, one might ask if the endpoint in 1970 could not be bent further upward by going outside of our stipulated interval of elasticities, $(1 - \rho)^{-1} \in [0.1, 0.8]$. 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 more than 10 percent from its historical average. (For a discussion of labor share predictions, see section 4.3.)

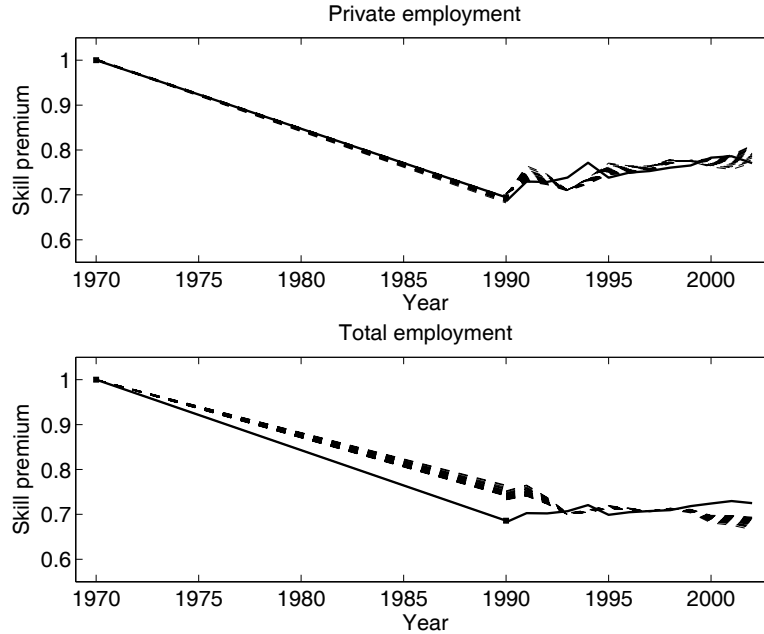


Figure 7: Swedish skill premium, given calibration period 1970-2002, in private employment (upper panel) and total employment (lower panel). The solid lines refer to data, and the dashed lines depict simulations.

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.¹⁸

¹⁸The 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.

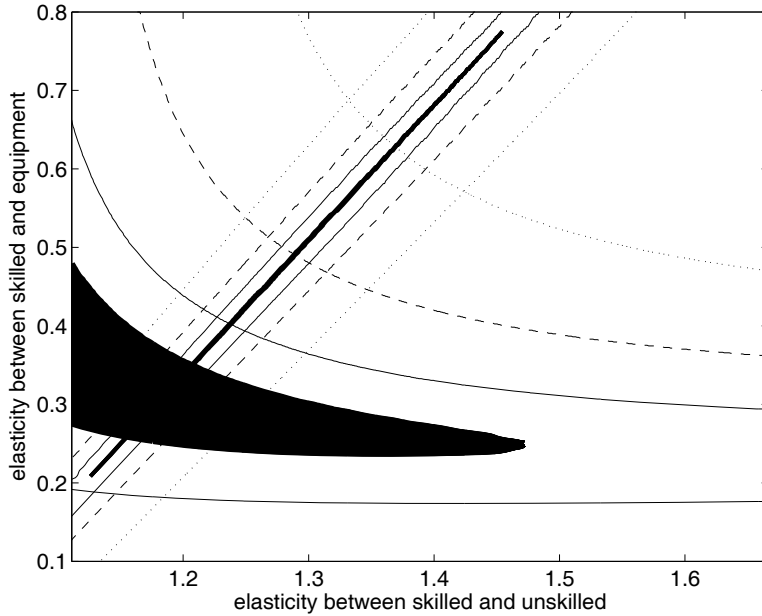


Figure 8: Contour maps of the coefficient of variation of the root mean square deviation with respect to the skill premium (contour lines with southwest-northeast orientation) and the labor share (contour lines with northwest-southeast orientation), in the space of elasticities for the Swedish model using private employment.

4.3 Best estimate of Swedish elasticities

Our analysis shows that total employment cannot explain the Swedish skill premium, but private sector employment can. To further refine our estimate of elasticities in private employment, we exploit observations on the labor share of income, and thereafter verify that other implications of the model are also consistent with data.

Recall that the dashed lines in the upper panel of figure 7 display simulations for elasticities of substitution between skilled labor and equipment that span the interval $[0.1, 0.8]$. As described in section 3.3, for each elasticity in that range, we have calculated a corresponding elasticity of substitution between skilled and unskilled labor so that the pair of elasticities minimizes the squared deviations between simulated and actual skill premia. Figure 8 shows a contour map of the coefficient of variation of the root mean square deviation in the space of elasticities, where the upward-sloping diagonal lines are the contour lines for the skill premium. The solid area, or better described as the thick line, in the middle of those contour lines depicts the set of elasticities that yields an average deviation of less than 2.5% from the actual skill premium. The solid thinner lines that surround the thick line are the contours of

an average deviation of less than 5%. Thereafter, the dashed (dotted) lines are the contours of an average deviation of less than 10% (20%). Thus, there is a deep valley in this contour map where by picking one elasticity, the other elasticity is pinned down fairly precisely. And since the thick line extends far in its southwest-northeast orientation, there are many pairs of elasticities that are associated with a good fit in the upper panel of figure 7. To further discriminate between these successful pairs of elasticities, we now turn to another dimension of inference.

For the simulation associated with each pair of elasticities, we compute the implied evolution of the labor share of income. As noted by KORV, the specification of the production function does not ensure that the labor share stays constant over time as it roughly does in the data. Ignoring the time variation in the actual labor share, the contour lines with northwest-southeast orientation in figure 8 depict deviations from the historical average of the labor share, i.e., our calibration target for the labor share. The configuration of contour lines is the same as that for the skill premium, i.e., deviations of less than 2.5%, 5%, 10% and 20%, respectively, but now with respect to the labor share. As compared to the skill-premium contour map, the labor-share contour map is considerably flatter. Taken together, the two contour maps suggest that the best estimate of Swedish elasticities lie in the southwest corner of figure 8. For example, if we want deviations of less than 2.5% with respect to both the skill premium and the labor share, the elasticity of substitution between skilled and unskilled labor falls in the range $[1.16, 1.22]$ with an associated range for the elasticity of substitution between skilled labor and equipment of $[0.26, 0.36]$. We will use the midpoints, an elasticity of substitution between skilled and unskilled labor (between skilled labor and equipment) equal to 1.19 (0.31), when we now turn to additional implications of our analysis.¹⁹

4.4 Additional implications

The first half of table 1 summarizes the country-specific calibration of our models for the U.S. and Sweden, which have successfully explained the skill premia over the period 1970-2002. Additional key implications of those simulations are reported in the second half of table 1,

¹⁹Our two-stage approach of arriving at the best estimate of Swedish elasticities in figure 8 is not only informative about the determinants but also nearly identical to a conventional calibration of the parameters simultaneously under a single loss function. Specifically, with equal weights on the percentage deviations of the skill premium and the labor share, the calibrated elasticity of substitution between skilled and unskilled labor (between skilled labor and equipment) would be 1.18 (0.31), with only minor changes to the other numbers reported in table 1. As suggested by figure 8, the calibration is fairly robust to varying the relative weights on the calibration targets of the skill premium and the labor share.

which lend further empirical support to our analysis.

Table 1: Calibration and simulations

	U.S.	Sweden
Country-specific calibration		
elasticity between skilled and unskilled labor	1.67	1.19
elasticity between skilled labor and equipment	0.67	0.31
capital structures share of income	0.117	0.170
labor share of income (base year 1970)	0.667	0.610
Simulated statistics 1970-2002		
average labor share of income	0.642	0.608
annual productivity gain ^a	2.0%	1.9%
annual growth rate of equipment (efficiency units)	7.8%	7.1%

^a Measured ‘per hour,’ i.e., we normalize $h_u + h_s = 1$ in each year.

The first implication in table 1 concerns the labor share of income. As discussed above, the specification of the production function does not guarantee a constant labor share. But, as in the original analysis of KORV, the model for the U.S. yields a stable labor share at roughly two-thirds of income throughout the period. The elasticities for Sweden were picked in section 4.3 to similarly yield a stable labor share, close to its historical average of 0.61.

The second implication concerns productivity growth. The simulated statistic in table 1 is computed under the assumption of no neutral technological growth, i.e., a time invariant parameter A in production function (1). Hence, we ask if productivity gains can be explained exclusively by equipment-specific technological change, as Greenwood et al. (1997) found to be the case for the U.S. in a similar framework but with a single type of homogenous labor.²⁰ Our simulations predict that output per hour grew by 2 percent per year in the U.S. and by 1.9 percent in Sweden, which are in line with what is observed in the data.²¹

²⁰According 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.

²¹Using data from the Groningen Growth and Development Center (<http://www.ggdc.net>, the Total Economy Database, ted07L.xls) the average growth rate in real GDP per hour between 1970 and 2002 was 1.9

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 7.8 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 7.1 percent. How does this fit the facts? While there exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon, Statistics Sweden reports data on capital stocks measured using similar methods as in the U.S. National Income and Product Accounts. KORV reports that U.S. equipment, as measured in the National Income and Product Accounts, grew at an annual rate of 3.4 percent between 1980-92. According to Statistics Sweden, the Swedish stock of equipment grew at a similar rate of 3.5 percent per year during the same period.²² While not conclusive, this suggests that the stocks of equipment developed similarly in both countries, as in our simulations.

5 Sweden through the lens of our model

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

5.1 Public sector expansion in the 1970s and 1980s

Public employment in Sweden grew to more than one third of total employment in the 1980s. Though, in our model it is not size *per se* that matters, but rather how public employment affects the ratio of skilled to unskilled labor in private employment. As mentioned in the introduction, a critical assertion by Lindbeck (1997, p. 1311) is that the public sector has particularly bolstered demand for labor at the lower skill spectrum. Indeed, the sharply increasing ratio of skilled to unskilled labor in private employment relative to that in total employment between 1970 and 1990 in figure 3, implies a sharply increasing ratio of unskilled to skilled labor in public employment. Using household survey data, Björklund and Freeman (1997, pp. 62-63) document an associated disproportionate increase from 1968 to 1991 in the

percent in Sweden and 1.6 percent in the U.S.

²²See Statistical report N 10 SM 9501 ("National Accounts 1980-1994": Appendix 3, "Stocks of fixed assets and national wealth 1980-1995").

fraction of workers in the lowest quartile in the hourly earnings distribution who are employed in the public sector. They emphasize that “[t]he pattern in most countries is for low-skilled workers to be underrepresented in the public sector, and finding equal representation [in 1991] is therefore surprising. . . . We interpret the upward trend and their proportionate representation in Sweden as signs that the public sector has, indeed, been a greater demand of their labor than in other countries.”

Given that essentially all Swedish employment growth in the 1970s and 1980s was in services provided by the local government, Rosen (1997, p. 80) concluded that “the welfare state encourages extra production of household goods and discourages production of material goods. From the normative view of economic efficiency, too many people provide paid household (family) services for other people, and too few are employed in the production of material goods. From the view of positive economic analysis, this is what explains the growth of local government employment of women and the growth of the welfare state.” Our analysis shows how the forces identified by Rosen and others have also had equilibrium repercussions on the Swedish skill premium.

5.2 Crisis and public sector stagnation in the 1990s

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

Through the lens of our model, none of these events would have had any effects on the Swedish skill premium unless they affected the evolution of the ratio of skilled to unskilled labor in private sector employment. After computing average annual growth rates of the ratio of skilled to unskilled labor in total employment and private employment, respectively, we conclude that there were indeed changes in this dimension between the periods 1970-1990 and 1990-2002. Specifically, while the annual growth rate of the ratio of skilled to unskilled labor in total employment slowed from 4.3% to 3.2%, the corresponding change in private employment was larger from 6.2% to 4.0%. Thus, the stagnating public sector in

the 1990s did not absorb unskilled labor to the same extent as in the previous two decades. Observers of the Swedish labor market who have focused on measures of total employment have not registered this detrimental development for unskilled labor. The theory informs us that unskilled workers are better off when the public sector increases its demand for their services, because the ensuing relative scarcity of unskilled workers in the private sector drives up their marginal product and wage. Our analysis attributes the reversal in the 1990s from a falling to an increasing Swedish skill premium to two factors: (1) the Swedish growth in the ratio of skilled to unskilled labor in total employment slowed after a couple of years into the 1990s, as shown in figure 2; and (2) the disproportionate hiring of unskilled labor in the Swedish public sector weakened in the 1990s, as reflected in figure 3. Next, we will quantify how much the Swedish skill premium was affected by the fact that the public sector has favored unskilled labor.

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

We use our calibrated Swedish model in table 1 to answer the following counterfactual question. Suppose that the relative growth rate of the ratio of skilled/unskilled labor in Swedish private employment as compared to this ratio for total employment had been the same as that of the U.S. over the period 1970-2002, what would have been the Swedish skill premium in such a scenario where the public sector had not favored unskilled labor? Specifically, we keep the actual time series for the ratio of skilled to unskilled labor in Swedish total employment, but let post-1970 changes in the ratio of skilled/unskilled labor in private employment relative to that of total employment be governed by the corresponding evolution in the U.S.²³

Figure 9 shows that the Swedish skill premium would have increased somewhat between 1970 and 1990 if it had not been for the expansion of public sector employment which disproportionately hired unskilled labor. Moreover, moving into the 1990s with a slowing growth in the ratio of skilled to unskilled labor in Swedish total employment (the upper line

²³The counterfactual ratio of skilled to unskilled labor in Swedish private employment in year t is computed as

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

and $h_{it}^{n,j}$ is the labor quantity of skill type i ($i = \text{u[unskilled]}$ or $i = \text{s[killed]}$) in employment n ($n = \text{priv[ate]}$ or $n = \text{tot[al]}$) in country j ($j = \text{Sweden}$ or $j = \text{US}$).

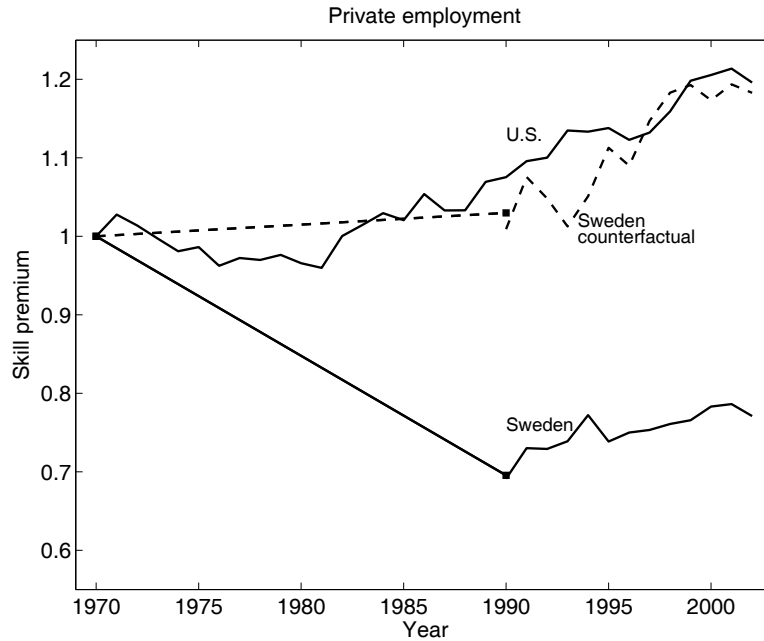


Figure 9: Skill premium in private employment in Sweden and the U.S. (normalized to one in 1970). The solid lines depict the data, while the dashed line refers to a counterfactual experiment for Sweden in which the evolution of the ratio of skilled/unskilled labor in private employment relative to that of total employment is assumed to be governed by the relative changes observed in the U.S.

in figure 2) coupled with an unchanging relative allocation of skilled and unskilled labor in private versus total employment (when assuming the U.S. evolution as given by the lower line in figure 3), our counterfactual scenario results in a Swedish skill premium in 2002 that is almost identical to that of the U.S.²⁴ Given this rather remarkable prediction, we conclude that public sector employment in Sweden has been critical for suppressing the Swedish skill premium over the period 1970-2002 as compared to the U.S. skill premium.

6 Discussion

Our thesis that public sector employment can explain the dramatic decline of the Swedish skill premium between 1970 and 1990 is a theory about both nonmarket and market forces. On the one hand, tax-financed local governments have disproportionately hired unskilled

²⁴The same answer to our counterfactual question is obtained when using Domeij and Ljungqvist's (2006) estimated Katz-Murphy (1992) equation for Sweden, where the log of the skill premium is a function of the log of the ratio of skilled to unskilled labor (in private employment), and a linear time trend.

workers in large numbers and thereby affected the ratio of skilled to unskilled labor available to the private sector. On the other hand, and in contrast to some other nonmarket theories, unskilled and skilled labor in our analysis are paid their marginal products in the private sector.

There are two ways that the allocation of labor between the public and the private sector could have arisen, and our analysis does not discriminate between them. First, suppose that trade unions and employer federations have set the observed wages. As long as the public sector stands ready to hire any surplus of workers at those wages (by acting as ‘an employer of last resort’), the observed labor quantities in the private sector are consistent with our theory, i.e., private employment is such that marginal products of unskilled and skilled labor are equal to the pre-determined wages. Second, suppose instead that the political process determines the size of the public sector and its hiring of workers. A well-functioning market economy would then ensure that the remaining labor quantities available to the private sector are employed at wages equal to the marginal products of unskilled and skilled labor.

Our theory cannot discriminate between these alternative scenarios because, like KORV, we take the quantities of unskilled and skilled labor as given. And like that earlier study, it is a shortcoming that we do not endogenize the countries’ skill formation and how such skill formation responds to a changing skill premium.²⁵ 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 finding is a prerequisite for any theory of a competitive private sector that maximizes profits, and can serve as a starting point for an extended framework with endogenous labor supplies.

Our analysis of the Swedish skill premium offers an important qualification to theories of how compressed wages outcomes in Europe are connected to the disproportional withdrawal of less skilled workers into nonemployment.²⁶ Relative wage outcomes do not only depend

²⁵It 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.

²⁶Expressing what he deemed to be the conventional wisdom, Krugman (1994, p. 51) explained high unemployment in many European countries as “an unintended byproduct of their redistributionist welfare states, and that the problem has worsened because the attempt to promote equality has collided with market forces that are increasingly pushing the other way.” Krugman ascribed those market forces to skill-biased technological change that reduces the earnings of unskilled workers relative to skilled workers, as in our KORV analysis, but where any withdrawal of unskilled workers from employment into benefit dependency mitigates those effects. A related argument is that European productivity gains are partly “artificial, as

on the relative employment of skilled and unskilled labor, but also how government policies affect the allocation of those who are employed. Albeit Sweden might be exceptional in terms of both the dramatic expansion of its public sector and the precipitous decline in its skill premium, our findings should give pause for a closer scrutiny of the interaction between public sector employment and relative wages.²⁷ Or more generally, how welfare state policies in Europe have affected the allocation of employed workers, and through that might have contributed to the suppression of skill premia in Europe.

Europe made labor expensive . . . firms were forced to slide northwest up their labor demand curves, retaining high productivity workers while forcing low-productivity workers into unemployment or out of the labor force entirely” (Gordon, 2006). For example, 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, as in our KORV analysis.

²⁷For evidence of considerable heterogeneity in public sector employment across the OECD countries, see Algan et al. (2002), and Gregory and Borland (1999).

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Appendix A Data Construction

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

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

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

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 63,593 in 1970, 88,864 in 1980,

80,620 in 1990 and 108,098 in 2002.

In the wage sample we follow Autor et al. (2008) by excluding (i) the self-employed, (ii) individuals with less than 40 weeks worked per year, (iii) individuals who work less than 35 hours per week, (iv) individuals with allocated income, (v) individuals whose weekly pay (after topcode adjustment, see below) is less than \$67 in 1982 dollars (using CPI as the deflator), or whose hourly wage exceeds 1/35th the topcoded value of weekly earnings.²⁸ The number of remaining observations are 32,897 in 1970, 41,958 in 1980, 44,385 in 1990 and 55,492 in 2000.

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

We classify individuals as private sector employees if their CPS industry classification belongs to the following set of CPS codes; (17-817, 849, and 888-899) for the years 1970-81, (10-411, 420-811, 841, and 882-893) for the years 1982-2001 and (111-6290, 6380-7790, 8560-9090, 9190 and 9290) for the year 2002. We have also used the Census of Population Survey’s classification into private and public sector employment with very similar results.

Prior to March 1989, wage and salary income is reported in the CPS as a single variable that is topcoded at values between \$50,000 and \$99,999. Beginning in 1989, wage and salary income is reported in two separate variables, corresponding to primary and secondary earnings, which are topcoded separately.²⁹ Like Autor et al. (2008), we handle topcoded earners as follows. We adjust primary and secondary earnings separately before summing them. For all years we multiply the topcoded values by 1.5. See Appendix A.4 for a sensitivity

²⁸We have investigated the following alternative specifications with unchanged results; (i) excluding individuals whose weekly earnings are less than 40 times half the hourly minimum wage, (ii) including individuals whose weekly earnings are less than 40 times half the hourly minimum wage but imputing their wages to equal half the minimum wage, or (iii) excluding individuals whose weekly earnings are less than 40 times one sixth of the hourly minimum wage.

²⁹Between 1989 and 1995 both primary and secondary earnings were topcoded at \$99,999. Between 1996 and 2002 primary and secondary earnings were topcoded at \$150,000 and \$25,000, respectively. In 2003 these values were raised to \$200,000 and \$35,000, respectively. Beginning in 1996, topcoded earners are assigned the mean of topcoded earners in the same sociodemographic group (sorted according to sex, race and worker status). We follow Autor et al. (2008) and reassign primary and secondary earnings to their topcoded values.

analysis with respect to different procedures of handling topcoded earners.

Appendix A.2 Swedish data

We construct the supply sample by excluding (i) individuals below age 16 and above age 64, (ii) individuals who are not in the labor force, and (iii) individuals without education or industry classification. The number of remaining observations are approximately 3.2 million in 1970, and between 3.6 and 3.9 million each year 1990-2002. We construct the wage sample by excluding the self-employed and those working less than 20 hours per week.

We classify educational attainment as in the U.S. data with one exception. Obtaining a traditional Swedish college degree requires 3-5 years of studies depending on field of specialization. We classify individuals who have completed a 3-year university or college degree as college graduates.

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.

Finally, when discussing how the composition of health care workers changed in the 1990s in the Introduction we use the following definitions. Health and elderly care is defined by SNI92 codes 85000-85330. Lacking information on professions we identify the least skilled worker category as those with a 2 or 3 year high school degree specializing in health and elderly care and we identify skilled staff as those with at least 3 years of university studies in medicine.³⁰

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

³⁰Corroborating statistics from the employers' association can be found in "Fakta om vård- och omsorgspersonal i kommuner och landsting", 2000, Swedish Association of Local Authorities and Regions. See graph 2 and page 5 for the least skilled and graph 12 for medical doctors.

hourly wage rate W_{gt} for group g in year t as

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

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

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

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

The groups are then aggregated into two classes; ‘skilled’ workers, G_s , and ‘unskilled’ workers, G_u . For each class $j \in \{s, u\}$ and year t , the class labor supply is computed as $h_{jt} = \sum_{g \in G_j} H_{gt} \bar{W}_g v_{gt}$, where the weights \bar{W}_g are the group wages 1990;³¹ 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 last year’ but prior to 1976 we only have data on ‘hours worked last week.’ Like Autor et al. (2008) we impute ‘usual hours worked last year’ as follows. Using data from the years 1976-1978 ‘usual hours worked last year’ is regressed on ‘hours worked last week’ (using the following set of dummy variables; 0, 15-29, 30-34, 35-39, 40, 41-48, 49-59 and ≥ 60 hours), a dummy variable indicating labor force status in the survey-year, a dummy variable indicating full-time status last year (i.e., working at least 35 hours per week) and interaction-terms between all these dummy variables and full-time status last year. The regression coefficients from this regression is then used to impute ‘usual hours worked last year’ for the years prior to 1976. 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 1976-1978.

Fourth, in the Swedish data for 1970 and 1990 hours worked is reported in intervals. For

³¹As a sensitivity analysis, we have tried three alternative weighting schemes; 1970 group wages, 2002 group wages, and average group wages across time, respectively. The results were similar.

1970 the intervals are 1-19, 20-34 and ≥ 35 hours per week, and for 1990 the intervals are 1-15, 16-19, 20-34, ≥ 35 . We have assigned hours per week using midpoints, assuming 43 hours per week for 1970 and 40 hours per week for 1990 in the upper interval.³² A potential problem with this procedure is that the implied average numbers of hours per week do not exactly match the ones reported by Statistics Sweden (see www.scb.se; and Pohjolassa, 1983). In the latter data, the average hours worked in 1970 (1990) was 39.3 (36.8) hours per week, with men working 43.5 (40.4) and women 32.8 (32.7) hours per week. Using midpoints implies that average hours worked in 1970 (1990) was 39.9 (36.4) hours per week, with men working 42.6 (38.7) and women 35.8 (33.9) hours per week. That is, using midpoints overestimates the working hours of women and underestimates the working hours of men. To investigate the importance of this mismatch we perform the following robustness check. The midpoints in each interval is multiplied by two factors; the first capturing aggregate time effects and second capturing time specific gender effects. 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 our 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 worked in each group g has remained constant at its 1990 level. This is roughly consistent with hours worked by age and gender as reported by Statistics Sweden (see www.scb.se). Between 1990 and 2000 average hours worked for employed men remained constant while it increased by 3 percent for employed women. In the age dimension, only a small fraction of the labor force experienced large changes; individuals age 19 and younger. To investigate the importance of the assumption of constant hours in each group g , we perform the following robustness check. Average hours in each group g is adjusted by a time specific age \times gender factor such that average hours worked by age \times gender groups are consistent with Statistics Sweden's estimate for each year in the period 1991 to 2002. The results of our paper were not affected.

Appendix A.4 Topcoding in U.S. data

As mentioned above, data on wage and salary income is topcoded in U.S. data. Figure 10 shows how the share of individuals who are topcoded in our wage sample has increased over

³²In the upper interval, we choose these hours to match the maximum number of regular working hours per week stipulated by law. While the 40 work-hour week was introduced in 1970, it was not fully implemented until 1973.

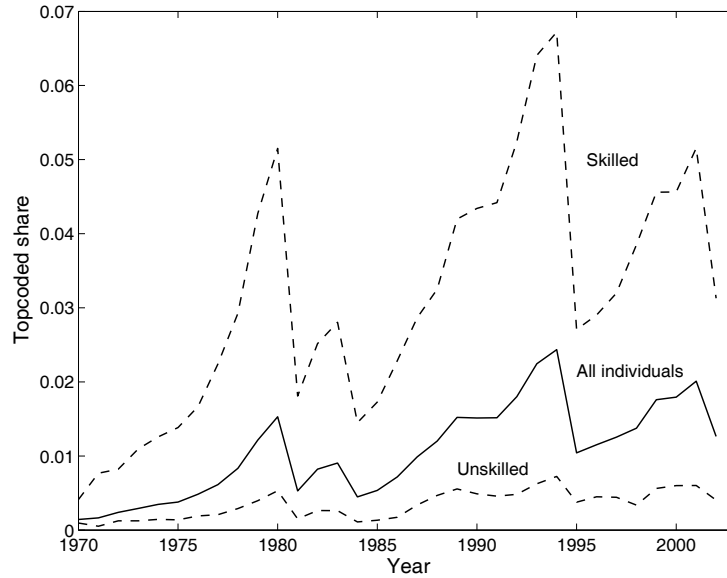


Figure 10: Share of individuals who are topcoded in U.S. data. The solid line depicts the topcoded share among all individuals while the upper (lower) dashed line shows the topcoded share among skilled (unskilled) labor. The topcode classification is based on total earnings until 1988, and primary earnings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

time (but whenever the topcoded value has been raised, there is a sharp drop). We therefore believe it is important to investigate the sensitivity of the U.S. skill premia to different procedures of handling topcoded earners. Besides our benchmark procedure of multiplying topcoded values by the adjustment factor 1.5, we consider three alternative procedures. In the first procedure, denoted NO, we follow KORV and keep the topcoded values without multiplying by any adjustment factor. In the second procedure, denoted CD, we follow Card and DiNardo (2002) who recensor primary and secondary earnings after 1989 to the 1988 topcoded values of \$99,999 and \$25,000, respectively, and then like KORV use topcoded values without multiplying by any adjustment factor. In the third procedure, denoted P, we assume that the top decile of the earnings distribution is Pareto distributed,³³ 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

³³For a recent study using the Pareto distribution to interpolate income distributions, see Piketty (2003).

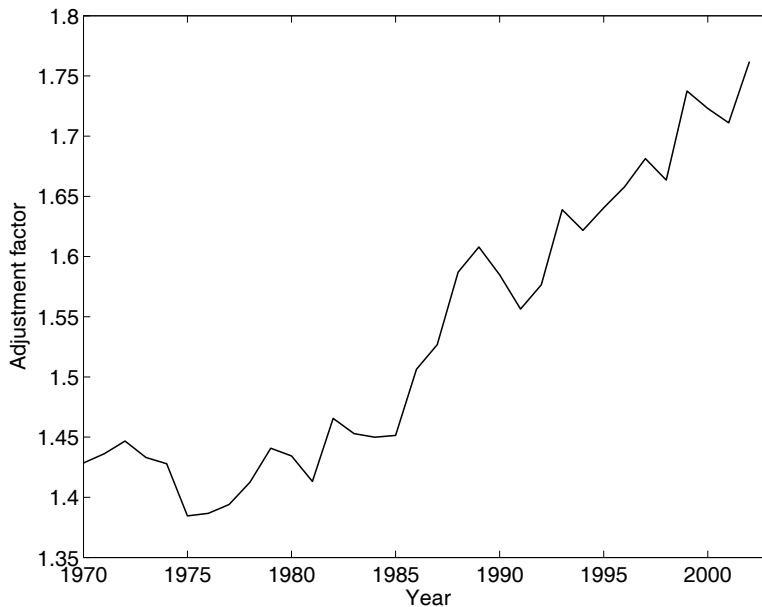


Figure 11: The adjustment factor $a_t/(a_t - 1)$ that topcoded earnings are multiplied by under the assumption that the top decile of the earnings distribution is Pareto distributed.

follows. Let $\vartheta_t(e)$ denote the fraction of earners with income greater than e in year t , i.e., $\vartheta_t(e) = (b_t/e)^{a_t}$, and after taking logarithms, $\log \vartheta_t(e) = \text{constant}_t - a_t \log e$. Based on this relationship between the fraction of earners with income above e and e , we compute OLS estimates of a_t . The data points $\{\vartheta_t(e_{it}), e_{it}\}$ are constructed from the earners in the top decile who are not topcoded. (The R^2 values in these regressions are above 0.98.) Figure 11 shows the implied annual estimates of the adjustment factor, $a_t/(a_t - 1)$, which are then used in our Pareto topcoding procedure. The adjustment factor is approximately 1.45 until 1985 and then gradually increases towards 1.7.

Figure 12 shows the U.S. skill premium in private employment for the benchmark and the three alternative procedures. The benchmark and the Pareto procedure yield very similar results. The NO procedure produces a skill premia that is always below the benchmark and the Pareto procedure. This is due to the fact that there are relatively more skilled individuals that are topcoded as shown in Figure 10. This is also why the CD procedure predicts that the increase in the skill premium was an episodic event of the 1980s without much change in the 1990s; there are relatively more skilled individuals that have their earnings recensored to their 1988 values.

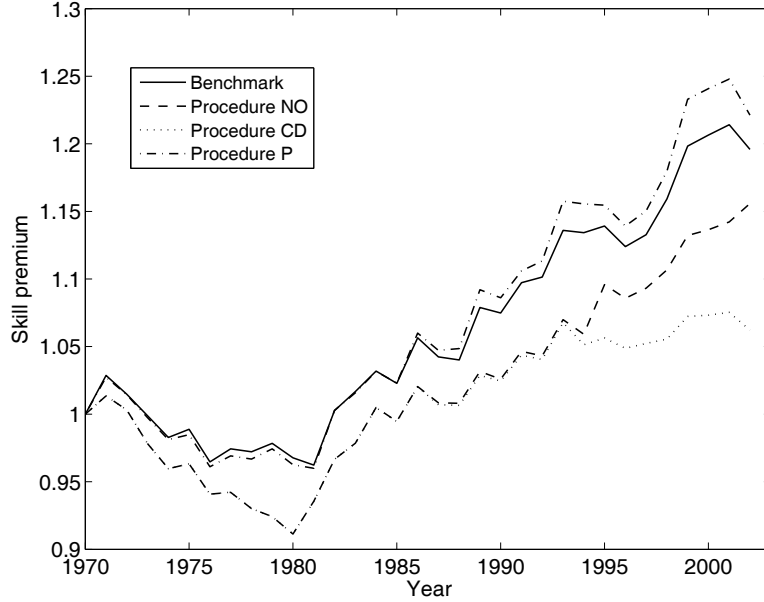


Figure 12: U.S. skill premium in private employment for various procedures to handle topcoded earners.

Appendix A.5 Income shares in Sweden

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

The income share for structures (α) is calculated as the annual average over the period 1970-1995, using stocks of structures and equipment and assuming that the rental rates on structures and equipment are given by $r_s = 0.09$, and $\tilde{r}_e = 0.209$, respectively, as implied by premises a)-c) in section 3.2. There exists no data on Swedish capital stocks in efficiency units similar to those reported by Gordon (1990), but noting that $r_e k_{e,t} = \tilde{r}_e \hat{k}_{e,t} \phi$ where $\hat{k}_{e,t}$ is the unadjusted stock of equipment at time t , $\phi = \left(1 - \frac{(1-\delta_e)}{1+\Delta}\right) / \left(1 - \frac{(1-\delta_e)p'_e/p_e}{1+\Delta}\right)$, and Δ is the average growth rate of investment in unadjusted equipment, we calculate the income share of structures as

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

The data on equipment and structures are from Statistics Sweden Statistical reports N1981:2:5

(Appendix 2, “Capital formation and stocks of fixed capital 1963-1980”) and N 10 SM 9501 (Appendix 3, “Stocks of fixed assets and national wealth 1980-1995”). The average growth rate of investment in equipment is 3.5 percent. We compute an income share of structures equal to 0.17 in both the private sector and in the total economy.

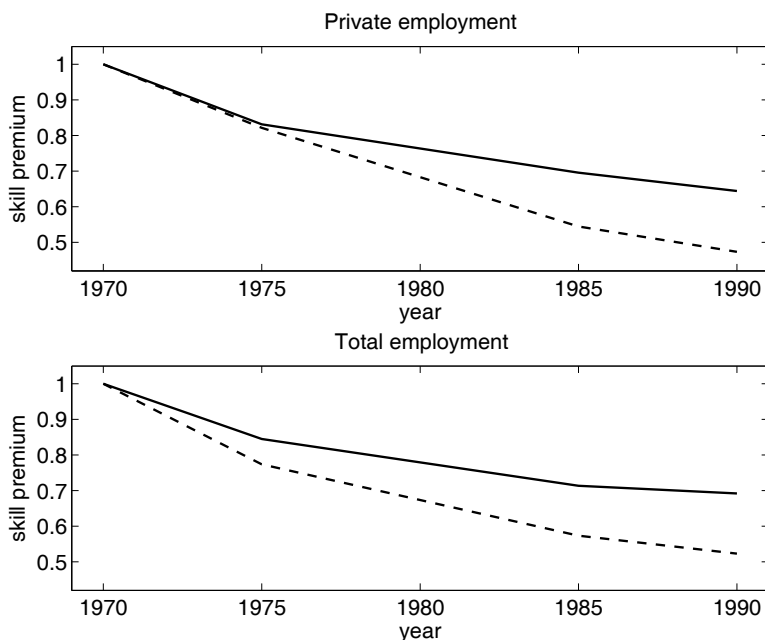


Figure 13: Swedish skill premium based on skilled versus unskilled occupations, in private employment (upper panel) and total employment (lower panel). Occupations are classified as skilled (unskilled) if at least a fraction x of workers in the occupation are skilled (unskilled) in both 1970 and 1990. The solid (dashed) lines refer to $x = 0.6$ ($x = 0.99$).

Appendix B Swedish Skill Premium 1970 vs. 1990

This appendix addresses two concerns regarding the dramatic decline in the Swedish skill premium between 1970 and 1990. The first concern is that the evidence rests on only two years of data. While Statistics Sweden carried out the Census of Population Surveys every five years up until 1990, only the censuses in 1970 and 1990 contain information on education. To lay this concern to rest we construct an alternative measure of the skill premium and show that it fell dramatically and continuously throughout the period. In particular, we proceed as follows. Every census between 1970 and 1990 include data on occupation for each individual. Using the observations on education in 1970 and 1990, we classify occupations as

being predominantly skilled (unskilled) if at least a fraction x of workers in the occupation are skilled (unskilled) in both 1970 and 1990, where we consider two alternative values for $x \in \{0.6, 0.99\}$.³⁴ The relative wage of skilled versus unskilled occupations is calculated using the procedure in Appendix A.3. Since the census in 1980 contains no information on earnings, we perform these calculations for the years 1970, 1975, 1985 and 1990.³⁵ As depicted in figure 13, the relative wage of skilled versus unskilled occupations fell by more than 30 percent, and the fall was a rather continuous event throughout the period. This is supporting evidence of a large drop in the skill premium between 1970 and 1990.

The second concern is that the computed skill premium might be due to less talented people getting a university education when the ranks of skilled workers expanded greatly, rather than reflecting a true drop in the relative wage of skilled versus unskilled labor. To lay this concern to rest we compute a skill premium only for workers who were employed in both 1970 and 1990. Excluding those whose educational classification changed between 1970 and 1990 leaves us with 1,219,011 individuals. We compute the relative wage of skilled versus unskilled for these workers using the procedure in Appendix A.3, and the weights that were calculated based on the entire wage sample. The resulting skill premium fell by 20 percent between 1970 and 1990 when using private employment, and by 27 percent when using total employment. Hence, workers employed already in 1970 experienced a similar evolution of their relative wage as workers who entered the labor force at a later date. Since Sweden between 1970 to 1990 experienced a large increase in female labor force participation, a related concern is that the results are driven by the selection of which women work. We have therefore computed the skill premium for men and women separately. We find that both when considering all workers and when considering those who were employed in both 1970 and 1990, men and women experienced a similar large decline in the skill premium.

³⁴With selection criterion $x = 0.6$, our procedure for classifying occupations as predominantly skilled or unskilled results in 90-98% of all workers in private and total employment, respectively, belonging to occupations that can be classified in census year 1970 and 1990, respectively. With the stricter criterion $x = 0.99$, fewer occupations can be classified and hence, the rates of worker coverage are smaller and lie in the range 48-62%.

³⁵The census in 1985 contains no information on hours worked, but as average hours worked in each subgroup of workers are almost identical in 1980 and 1990, we impute the 1985 values to be the average of those in 1980 and 1990.

Appendix C KORV's Model

Appendix C.1 Equilibrium properties

By solving for k_s from first-order condition (4) and substituting into production function (3), we get an equilibrium expression for output

$$y = (A\Omega)^{\frac{1}{1-\alpha}} \left[\frac{\alpha}{r_s} \right]^{\frac{\alpha}{1-\alpha}} \left\{ \theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}} \right\}^{\frac{1}{\sigma}}. \quad (11)$$

From first-order conditions (5) and (7), we get an equilibrium expression for the quantity of equipment,

$$k_e = \left[\frac{w_s}{\theta_s r_e} \right]^{\frac{1}{1-\rho}} h_s. \quad (12)$$

From first-order conditions (6) and (7), we get an equilibrium expression for the labor share of income, χ ,

$$\begin{aligned} \chi &\equiv \frac{w_u h_u + w_s h_s}{A F(k_s, k_e, h_u, h_s)} = (1 - \alpha) \frac{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma-\rho}{\rho}} \theta_s h_s^\rho}{\theta_u h_u^\sigma + (k_e^\rho + \theta_s h_s^\rho)^{\frac{\sigma}{\rho}}} \\ &= (1 - \alpha) \frac{1 + \pi \frac{h_s}{h_u}}{1 + \left[\pi^\sigma \theta_u^\rho \theta_s^{-\sigma} \left(\frac{h_s}{h_u} \right)^{\sigma(1-\rho)} \right]^{\frac{1}{\sigma-\rho}}}, \end{aligned} \quad (13)$$

where $\pi \equiv w_s/w_u$ is the skill premium, and the last equality in (13) is obtained by twice invoking expression (10).

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

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

This claim can be verified as follows. First, after invoking equilibrium expressions (11) and (12) for output and equipment, we confirm that first-order condition (7) for the em-

ployment 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 (6) 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 (13) is also unchanged, i.e., we verify that $\hat{\theta}_u^\rho \hat{\theta}_s^{-\sigma} = \theta_u^\rho \theta_s^{-\sigma}$.

Appendix C.2 Calibration to the base year

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

1. For an arbitrary value of Ω , we can compute the labor share of income as a function of the parameters θ_u and θ_s . Such a mapping is illustrated in figure 14 for one particular value of Ω . Given a calibration value for the labor share, say $\chi = 2/3$, the bold curve in figure 14 depicts pairs of parameter values (θ_u, θ_s) that produce the target value χ .
2. Both the labor share of income and the skill premium are constant along the bold curve in figure 14, as described in Claim 1. Hence, given the calibration value of the labor share, the bold curve in figure 14 represents a mapping from the parameter Ω (that is held fixed in the graph) to a specific value of the skill premium, π . Given the calibration value of the labor share, the compilation of pairs (Ω, π) from successive graphs when varying Ω , yields a mapping as illustrated in figure 15.
3. Given a calibration value for the skill premium, figure 15 pins down our choice of parameter Ω . Given that value of Ω , the associated figure 14 contains a bold curve depicting permissible parameter values θ_u and θ_s . On the basis of Claim 1, we can find parameter values (θ_u, θ_s) to match any absolute wage level; we set $w_s = 1$.

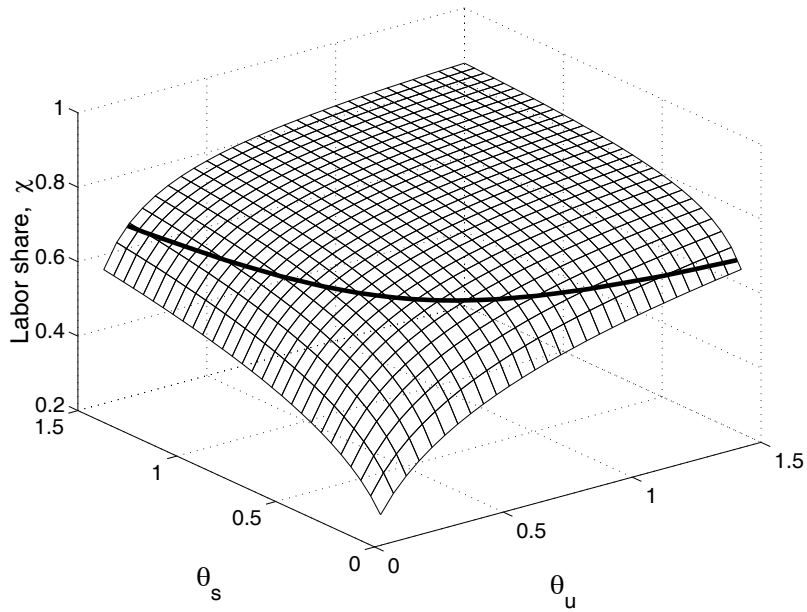


Figure 14: Labor share of income, χ , as a function of parameters θ_u and θ_s , for a given value of the parameter Ω . Along the bold curve is not only the labor share constant, $\chi = 2/3$, but so is the skill premium.

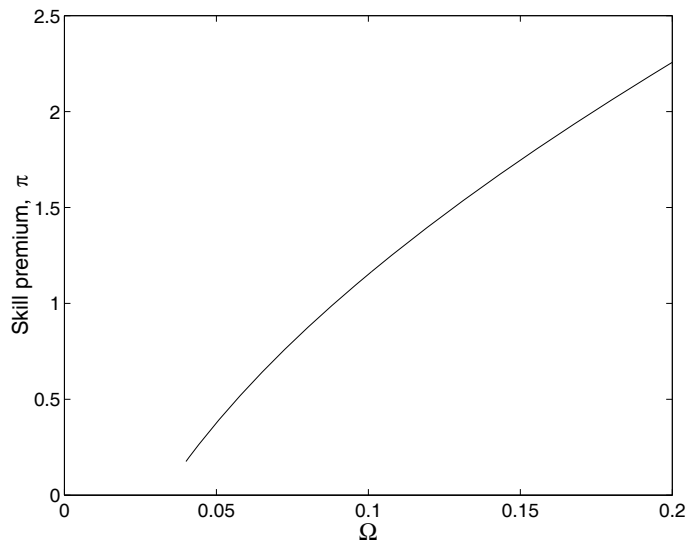


Figure 15: Skill premium, π , as a function of the parameter Ω , given that the labor share of income is held constant, $\chi = 2/3$.