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Individual Wage Setting, Efficiency Wages and Productivity in Sweden*

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Abstract

Swedish wage setting has undergone drastic changes during the last 10-15 years. While Sweden was known for its narrow wage distribution, wage differentiation and wage bargaining at the individual level has become leading principles among white-collar workers' unions. The purpose of the present paper is to analyse the consequences of this wage policy shift. Wage differences have increased drastically among white-collar workers while remained constant or even decreased among blue-collar workers. We show that wage differentiation has had a strong effect on white-collar workers' *average* wage, and caused a major increase in the wage gap between the aggregates of white-collar and blue-collar workers. We also show that increases in the coefficient of variation of wages raise productivity in firms with many workers in that worker category. Last and foremost, we show that the transition to individual wage setting raises the scope for firms to set efficiency wages and we find support for the fair wage version of efficiency wage setting. The effects of higher wage/fair-wage rates were stronger in the late 1990s, when wage differentiation increased more, than in the early 2000s.

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Introduction

For many years, the Swedish labour market was characterised by centralised wage setting combined with very narrow wage differences in an international comparison. This was the result of the “solidaristic” wage policy that was a leading principle particularly among blue-collar workers’ trade unions. On top of that, the main political objective of the ruling Social democratic party was one of economic equality, manifested in a highly re-distributive fiscal policy. The end result of this corporative policy was that the after-tax income differences became extremely small in Sweden when compared to those of other countries.

Wage compression reached a peak in the first part of the 1980:s. Since then, and particular during the last 10 to 15 years, Sweden has experienced major changes in wage setting practices and in wage distribution. First, there has been a clear trend of decentralisation towards individual wage bargaining, particularly so among white-collar workers. Secondly, this decentralisation has been followed by drastic increases in wage differences among white-collar workers. This development is in sharp contrast to the situation among blue-collar workers for which unions have maintained very narrow wage differences, though the scope for individual wage setting has increased also among these groups. Another prominent feature of wages in Sweden is a considerable widening of wage differences across the aggregates of blue- and white-collar workers.

In this paper we analyse the consequences of the transition to individual wage setting and to a policy of wage differentiation that characterized the white-collar unions after the mid 1990:s. These changes can be expected to raise the scope for efficiency wage setting among white-collar workers and we shall test to what extent this has been the case. Much research has focused on the distribution of wages within and between firms.[1] One problem with that approach from the point of view of efficiency wage setting is that wage comparisons to a large extent are between non-peers, which are not likely to have any major effects on work morale or incentives to perform well. To analyse the effects of comparisons with peers, we therefore analyse the effects of wage differentials within worker categories on firms’ productivity. Basically, this test amounts to examining if the prospects of better pay in worker categories having large wage differences, is conducive to firms’ productivity. Using a large worker-firm data set, we find that firms having a large number of workers in categories where wage differences are large perform better in terms of labour productivity and the

[1] See e.g. Levine (1991), Akerlof and Yellen (1988) and Hibbs and Locking (2000).

change in wage policy has contributed significantly to raise productivity within firms. This suggests that workers' effort is stimulated by the prospects of higher pay. These prospects for higher incomes may come as a result of the individual signalling strong work effort to the current employer or to potential outside employers.

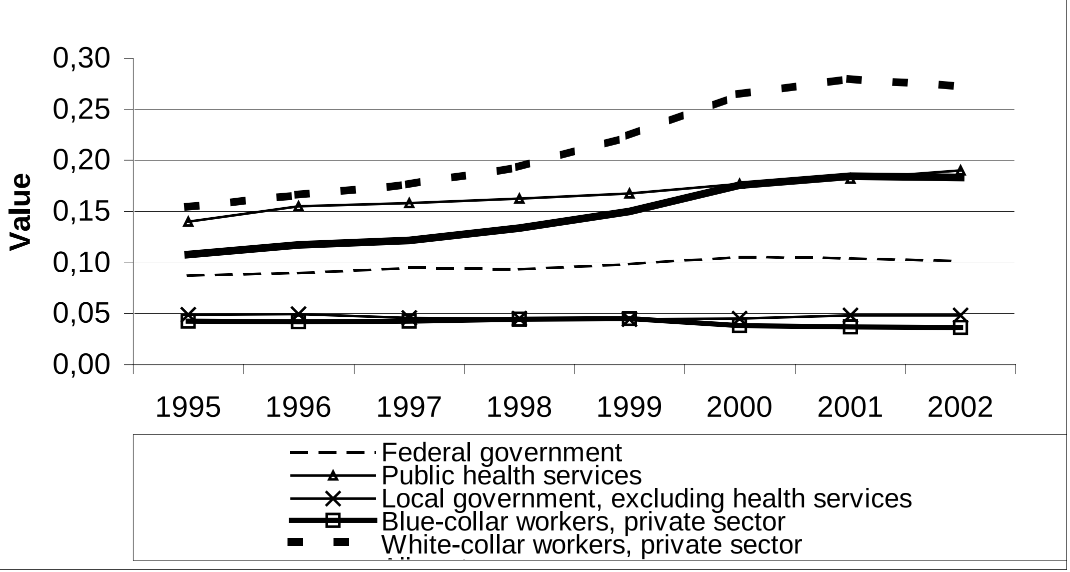
Apparently, the decentralisation of Swedish wage setting has increased the scope for efficiency wage setting. Is the increase in wage differences an expression for efficiency wage setting? We extend the analysis by studying the effects of fair wage setting on productivity. We argue that the perceived fair wage can be predicted from estimated Mincer equations that include, besides individual characteristics, also sector variables and the firm's financial result per capital invested. We find that firms with many employees that receive a wage that is high compared to their perceived fair wage, are more productive than other firms. This supports the fair-wage version of efficiency wage setting and our empirical results suggest that much of the boost of Swedish productivity in the 1990s is due to changes in wage setting practices.

Trends in Swedish wage formation

After the peak in wage compression in the mid 1980s a period of wage differentiation followed among white-collar workers.[2] As seen in *Figure 1*, wage differences, here measured by squared coefficients of variation, have continued to increase and accelerated since the mid 1990:s in most sectors reported. Towards the end of the 1990:s, wage differences increased at a remarkable rate among white-collar workers in private sectors.[3] Notable is also that the coefficient has *not* increased and even decreased slightly among blue-collar workers in the private sector.[4] We also see large increases in wage dispersion among white-collar workers in the public health sector but not among workers in local government, belonging to blue-collar workers' unions, outside the health sector.

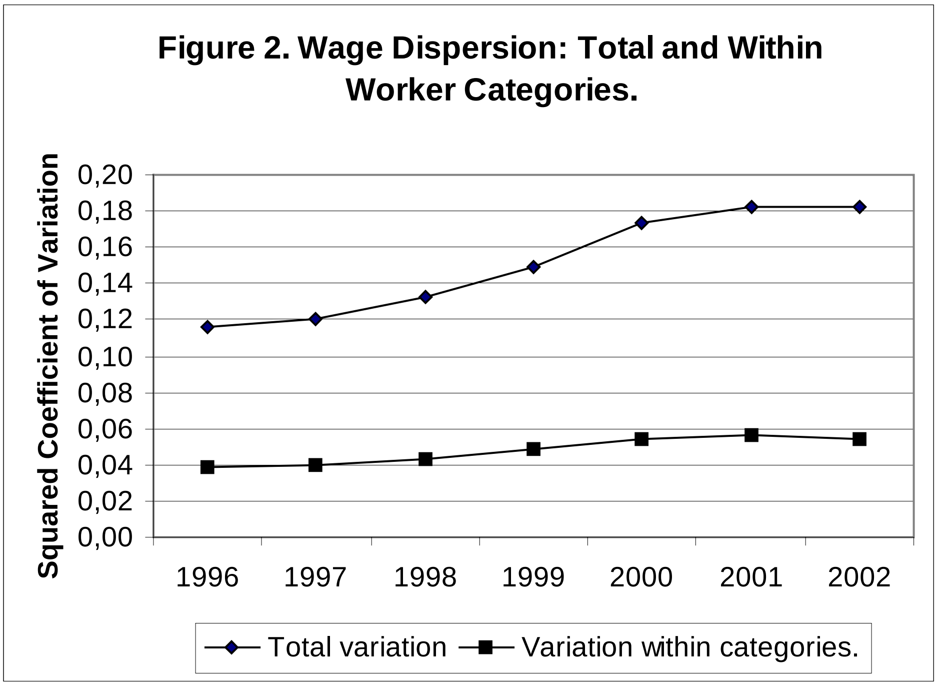
[2] The fairly modest increases in wage differences since the mid 1980:s have been documented in e.g. OECD (1996).
[3] These are workers in categories covered by the two white-collar worker confederations TCO and SACO.
[4] Workers in categories covered by blue-collar workers' confederation, LO.

Figure 1. Squared coefficients of variation. 1995-2002



Source: Calculated from FIEFs data base.

Increasing wage differentials among white-collar workers have driven up the economy wide variation of wages, which, as measured by the squared coefficient of variation, has increased from less than .12 in 1995 to more than .18 in 2002. It is of interest to see how much of this increase has taken place *within* the worker categories and how much *across* the worker categories. In *Figure 2* we show total wage dispersion and wage dispersion within the worker categories 1996-2002, the difference between these two curves is the variation of wages across worker categories. Wage dispersion has increased both within the worker categories as well as across the worker categories. The share of “within category” variation of the total variation has fallen slightly during the period, from .33 to .30. The large increases took place in the late 1990.s while wage dispersion was fairly stable in the early 2000’s.



Source: FIEFs data base.

Changes in relative wages: White- vs blue-collar workers.

We have documented substantial changes in the distribution of wages *within* the group of white-collar workers while the distribution of wages among blue-collar workers has remained relatively stable. A crucial issue for the robustness of Swedish wage formation is if the *average* wage of white-collar workers has increased compared to the *average* wage of the blue-collar workers.

Table 1 shows the average annual wage increases during the 1994-2003 period. The average increase during the period 1994 to 2003 was 44 percent for all workers, 47 percent for white-collar workers and 40 percent for blue-collar workers. In general, and for both groups of workers, wages increased more in the private sector than in the public. These differences in annual wage increases accumulate to large absolute wage differences between white- and blue-collar workers today.

Table 1. Average annual increases of monthly wages 1994-2003.

	Blue-Collar Workers	White- Collar Workers	All Workers
All sectors	4.44	5.22	4.89
Private sector	4.56	5.33	5.00
Public sector	4.00	4.89	4.56

Source:www.lo.se/home/lo/home.nsf/unidView/35CDC85CCAC75747C1256F17002CB995/$file/loner2004.pdf

A relevant objection to such industry sector data is that structural changes, manifested in workers changing to better paid jobs, have taken place within firms and these changes could affect the data on wage increases. This has been stressed as a possible bias, particularly for the white-collar workers, and data such as those in Table 1 could therefore be misleading. To investigate the importance of the structural effects we compare wage increases of the individual white-collar *worker categories* with those of blue-collar *worker categories* rather than the increases within sectors. *Figure 3* shows wage increases within 71 white-collar worker and blue-collar worker categories. Data are here limited to 1998-2002 since it is only for this period that we obtain comparable data series.

It is clear from this figure that also when we account for the structural effects by focusing on wage increases among well-specified worker categories, white-collar workers' wages have increased considerably more than blue-collar workers' wages. Measured as the arithmetic mean, white-collar workers' wages increased by 17.6 percent (or 4.4 percent per year) during 1998-2002 while blue-collar workers' wages increased by only 12.3 percent (3.1 percent per year).[5] This implies that when we correct for structural effects, the differences between white-collar and blue-collar workers' wages are even larger than those in Table 1, suggesting stronger structural effects among blue-collar workers than among white-collar workers.[6]

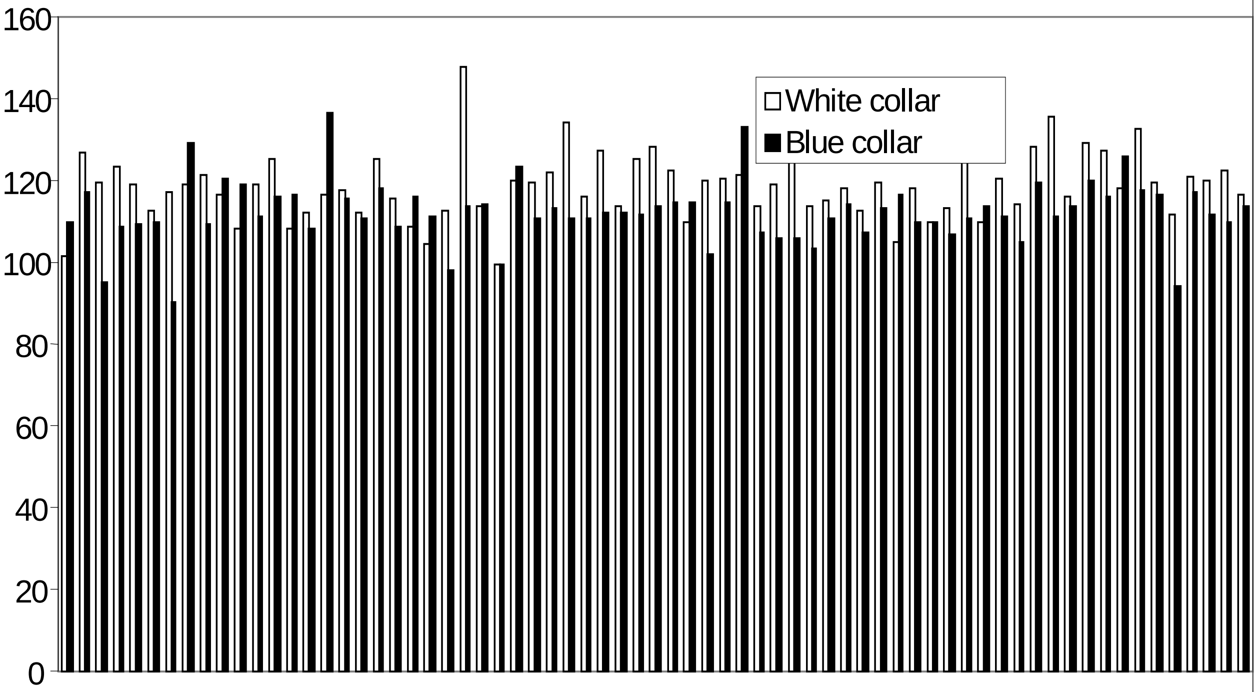
[5] Since we take the arithmetic mean we have excluded those worker categories that in our data set contain less than 100 observations so as not to exaggerate the weight of small worker categories.

[6] Note though that periods differ because of data limitations. The data on sectors are based on the period 1994-2003 while data on worker categories are based on the period 1998-2002.

Do increasing wage differentials cause wage drift?

Is there a connection between the considerably larger wage increases for white-collar workers reported in Tables 1 and Figure 3, and the increases in wage distribution among white-collar workers reported in for instance Figure 1? A possible explanation could be that since wages are upward flexible but downward rigid, a policy to raise wage differentials would cause a bias that raises also the *average* wage. It could also simply be viewed as a consequence of wage decentralisation that makes it possible to select workers that are underpaid compared to their productivity. These workers would then enjoy wage hikes while employers would find it hard to lower wages of workers that firms consider to be overpaid compared to their productivity. The latter group could find a necessary support in collective agreements that prevent a drop in wages.

Figure 3. Wage increases among worker categories 1998-2002. Index: 1998=100



Finally, the transition to individualised wage setting could imply that the scope for efficiency wage setting has increased, suggesting that firms would hesitate to reduce wages, which could reduce work effort but be inclined to raise wages so as to stimulate work effort. Irrespective of the reason, we shall refer to wage increases higher than those stipulated in the collective agreements as wage drift.

Before we proceed to the empirical analysis, we shall make a few clarifications. We shall argue that one should expect *any* change in the wage distribution, be it an increase or a decrease, to tend to raise the average wage under wage rigidity. Assume for simplicity two workers in the same worker category, one having a high wage w^h and one having a low wage w^l. Both wages are downward sticky. The average wage in this worker category is $w = (w^h + w^l)/2$. Now assume a union with a policy that allows the firm to increase the wage difference between workers, which can be accomplished by raising the wage of the high wage earner and/or lower it of the low wage earner. However, if wages are downward sticky, the effect on the average wage will be $dw = \frac{1}{2}dw^h + \frac{1}{2}dw^l = \frac{1}{2}dw^h > 0$ where $dw^l = 0$ due to downward wage rigidity. Since the firm cannot lower the wage w^l as a part of the strategy to increase the wage distribution, the average wage must rise.

However, we also see that a strategy to *reduce* the wage differences would have a similar effect. To reduce the wage distribution we would like to lower the wage of the high wage earner and to raise it for the low wage earner. In this case we would find that $dw = \frac{1}{2}dw^h + \frac{1}{2}dw^l = \frac{1}{2}dw^l > 0$

where $dw^h = 0$ since downward wage rigidity now applies to the high wage earner. Also in this case we would find that the average wage rises. Hence, any measure to affect the wage distribution, be it an increase or a decrease, will tend to raise the average wage.

To test if wage differentiation has caused wage drift, we first estimate a standard Mincer wage equations to which we add a variable to capture the effects of a shift in the policy of wage differentiation and individualized wage setting. This shift took place during 1996 and 2002. The question arises how to capture wage policy shifts.[7] A straightforward approach is to

[7] Penetrating the wage agreement could be highly misleading, besides resource demanding, since a wage agreement may *allow* for wage differentiation and individual wage setting though this need not be implemented by the parties. A case in point is the LO-union "Svenska kommunalarbetareförbundet" organising blue collar workers in the local governments. This union opened for the possibility of individual wage setting and wage differentiation in much the same way as the white-collar unions. However, at the local level, the representatives of the Svenska kommunalarbetareförbundet have not

consider the wage distributional outcome of the policy shift which had stabilized by 2002. To identify the worker categories for which the wage policy during the 1990:s has been one of wage differentiation, we specify the following continuous variable:

$$WD_j = \ln\left\{1+\left[wdiff_{2002,j} - wdiff_{1996,j}\right]/\,wdiff_{1996,j}\right\} \tag{1}$$

where WD_j is the wage policy as represented by the change in wage differences of worker category *j*, $wdiff_{t,j}$, and which we measure by either the squared coefficient of variation or percentile ratios. For each category of workers, the variable captures changes in wage distribution between 1996 and 2002. It is implicitly assumed that the outcome is in accordance with the objectives of the unions: As wage bargaining to a large extent is handed over to the individuals and to the firm's representative, the union agrees to the outcome which, as shown above, implies heavily increasing wage differentials among white-collar workers.

An alternative is to specify a dummy variable as:

$$WD_j^d = \begin{cases} 1 if\, 1+\left[wdiff_{2002,j} - wdiff_{1996,j}\right]/\,wdiff_{1996,j} > 1.05 \\ 0 if\, 1+\left[wdiff_{2002,j} - wdiff_{1996,j}\right]/\,wdiff_{1996,j} <= 1.05 \end{cases} \tag{2}$$

which takes on a unity value if wage differentials have increased by more than 5 percent, else a zero value.[8] The remaining variables are standard Mincer type variables like educational levels (compulsory schooling only is the basis), seniority measured as number of years with the present employer,[9] seniority squared, age, age squared and gender. We have a sample of 2 305 534 observations[10] for year 2002 distributed over 114 white-and blue-collar worker categories (professions).[11] For the *wdiff* variables there is, of course, only variation over these 114 observations. See Appendix for a description of data.

made use of this possibility, which is manifested in a more or less unchanged distribution of wages among the union's members.

[8] Five percent is chose arbitrarily but can, of course, easily be altered.

[9] For employees in the public sector, this variable is available only two years back.

[10] In the data set there are 2 261 514 individuals of which 65 439 have more than one job.

[11] Three-digit level.

We shall run the following regression:

$$w_{i,02} = \alpha + \beta X_{02} + \gamma WD_{i,02-96} + \varepsilon_{02} \qquad (3)$$

where X_{02} is a vector representing the standard variables in the Mincer equation, $WD_{i,02-96}$ is the change in the distribution of wages of the worker category to which worker i belongs and ε_{02} is the error term.[12]

In *Table 2*, *Model 1* shows the results without any wage distribution variable and the results are much in line with our expectations. *Model 2* shows the results when adding the dummy variable in equation (2), based on the 90/10 percentile ratios, PR_j. We see that the inclusion of the variable raises the explanatory power of the model as measured by the R^2 and that the estimated coefficient is significant.[13] *Model 3* and *Model 4* show the results when alternatives of the continuous variable in equation (1) are used. Adding the continuous wage distribution variable increases the explanatory power further compared to the results in *Model 2*. That the variable WD_j^d comes out significant and positive indicates that there is a positive effect on overall wages of a wage setting policy, as pursued by a large number of white-collar unions, that increases the wage differences among the employees.

In the regression denoted *Model 3*, the continuous variable WD_j in equation (1) is represented by the changes in 90/10 percentile ratios, PR_j. The coefficient quantifies the effect of the wage differentiation policy. The estimated elasticity is .7, implying that if a union raises wage differences so that PR_j increases by 1 percent, then the average monthly wage of that group rises by .7 percent. Consider, for instance, the effects of an increase in wage differentials among the aggregate of white-collar workers. Between 1996 and 2002, the variable P_{90}/P_{10} increased from 2.14 to 2.42 implying that our variable PR_j increased by 13 percent. The estimated

[12] We have a cross section variable for a single year, 2002, on the right hand side and we make no assumptions concerning the dynamics involved. Thus, by estimating equation (3) we test if the shift in wage policies that took place in many unions between 1996 and 2002, as measured by the shift in wage distribution, affected the wages of year 2002. In principle though, we expect the wage policies of 1996-2002 to affect also wages in the years following 2002. Since wage rigidity prevents wages from falling, the effects of wage policy shifts can be expected to be of a long run nature.

[13] We set the limit for a worker category that has pursued a wage differentiation policy somewhat arbitrarily at 5 percent. Alternatively, we tried a 10 percent limit, which also came out highly significant, but the overall explanatory power of the model was in this case lower.

value, .7 then implies a general wage effect of .7*13=9.1 percent for white-collar workers. Thus, *ceteris paribus*, more than 9 percentage points of the 47 percent increase in white-collar workers wages during 1994-2003 could be ascribed the policy of wage differentiation.

As documented in Figure 1, wage differentials among blue-collar workers have not increased and even fallen slightly during the same period. For worker categories with unchanged wage distributions, the narrowed wage distribution would lead to a drop in wages amounting to .14 percent. However, there is reason to question this latter result from a theoretical point of view. We argued above that efforts to raise wage differences are likely to add to a general increase and this effect has been strongly supported above. The reason is that wages are flexible upwards but rigid downwards. A *lowering* of wage differences cannot, however, be argued to *lower* wages in the corresponding way. To test for this, we performed a regression in which the distribution variable (1) was not allowed to take on a negative value but assumed a zero value for a worker category experiencing a drop in the distribution of wages.

The results are shown in *Model 4*, and the increased R^2 indicates that this formulation of the variable adds to the explanatory power. We also find that the estimated coefficient now is higher, close to 1.00. This implies that of the 47 percent wage increases during 1994-2003, 13 percentage points could be ascribed the wage differentiation policy. *Ceteris paribus*, had the white-collar unions pursued the same policy as blue-collar unions, average white-collar workers' wages would have fallen relative blue-collar workers' wages.

Finally, in *Model 5*, we estimate the model using the change of the log of the squared coefficient of variation, denoted *CofV*.[14] The squared coefficient of variation was used in *Figure 1* to show the changes in the distribution of wages. The advantage of this variable is that it gives a better representation of the changes along the full range of the wage distribution while the variables used above only compare the wage of the 90:th percentile to that of the 10:th percentile.

The explanatory power remains high. The estimated coefficient of $CofV_j$ is .21. Since the squared coefficient of variation of white collar workers increased by 80 percent (Figure 1) between 1996 and 2002, this estimate implies that 16 percentage points (80*.21) of the 47 percent total wage

[14] To obtain this variable we substitute the squared coefficient of variation for the ratio of the 90:th percentile to the 10:th percentile in equation (1). We also set the negative values of the variable to zero, i.e. as in Model 4 we do not allow for a lowered wage differences to have a negative effect on the wage.

increase can be ascribed the wage differentiation policy. This figure should be compared to the one obtained in *Model 4* (13 percentage points).[15]

Table 2. Regression results for 2002. Mincer wage equations with effects of shifts in wage policies.

	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**
High School 2 years	.0458 (50.60)	.0407 (45.57)	.04115 (46.25)	.0423 (47.74)	.0380 (42.86)
High School 3 years	.0927 (121.23)	.0844 (111.95)	.08273 (109.96)	.0869 (116.17)	.0804 (107.31)
University <3 years	.2644 (263.76)	.2380 (239.65)	.2374 (239.82)	.2412 (245.14)	.2337 (236.98)
University $\geq$3 years	.3295 (420.41)	.3221 (416.59)	.3140 (406.48)	.3206 (417.59)	.3066 (397.75)
Ph.D.	.6222 (411.90)	.5924 (396.76)	.5879 (394.58)	.5905 (398.52)	.6166 (416.98)
Age	.0206 (205.99)	.0210 (213.47)	.0210 (213.88)	.0210 (214.39)	.0208 (212.22)
Age^2	-.0002 (-162.69)	-.0002 (-171.12)	-.0002 (-171.78)	-.0002 (-171.55)	-.0002 (-169.42)
Seniority	.0104 (35.37)	.0118 (40.64)	.01115 (38.52)	.01153 (40.05)	.0118 (41.10)
Seniority^2	.0025 (53.25)	.0021 (45.25)	.0026 (55.59)	.0023 (49.62)	-.0021 (-45.52)
Gender	-.1685 (-517.45)	-.1645 (-512.06)	-.1697 (-530.14)	-.1617 (-506.38)	-.1731 (-542.77)
$WD_j=PR^d$		.0867 (260.78)			
$WD_j=PR_j$			.7025 (283.23)	.9957 (318.56)	
$WD_j=CofV_j$					.2124 (317.10)
No of obs.	2305534	2305534	2305534	2305534	2305534
Adj R^2	.3478	0.3665	0.3697	0.3753	.3750

Notes: Dependent variable is the log of wages in month-equivalents. t-ratios in parenthesis. Robust estimates. PR_j means that wage differences are defined from percentile ratios and $CofV_j$ from the coefficient of variation.

[15] As noted above, the wage policy variable only varies over the 114 worker categories while the other variables vary over the 2305534 observations. We also ran a regression with the mean wage of the category as the dependent variable and the wage distribution as in Model 5 as the explanatory variable. This yielded an estimated parameter of .22, i.e. very close to the estimated parameter in Model 5, .21.

Above, we found that the shift in wage policies between 1996 and 2002 affected the structure of wages in 2002. While we know that also low wage white-collar unions shifted to a policy of wage differentiation, it would still be relevant to ask if it is workers in the better paid categories that shifted to the policy of wage differentiation. This amounts to asking if it was the case that already in 1996 we had a similar wage structure as the one in 2002. To test for this, we may focus on the effects of the shift of wage policy on the individuals' wage *increases* between these years. A further advantage is that we may make use of the information on wages also in the first year, 1996. The results are presented in Appendix 2 and show that the results from these regressions on rate-of-change form generally support those on static form, Models 2-6. A policy of increased wage differentials, like that among white-collar worker unions, raises wage growth.

Productivity and the prospects of better pay

It is possible that increased wages among white-collar workers reflect an increased scope for firms to stimulate productivity increases that correspond to the wage increases. This would suggest that productivity among the worker categories that have been exposed to increased wage differentials has increased. It would also imply that wages have been set above the market clearing level for these worker categories, which would worsen workers' labour market situation. Finally, it could imply that increased wage differentials between white- and blue-collar workers is an adjustment to the market forces and that these differences adequately reflect differences in labour productivity between the groups. If this is the case, the new wage policy has raised the efficiency of the Swedish economy, though at the cost of reduced equality.

Crucial to our interpretation of the results is if the increased wage dispersion has contributed to raise productivity. We have no information about the individuals' productivity but with our worker-firm linked data set we can shed light on the effects of wage differentials within worker categories on firms' productivity. In this section we test the hypothesis that firms with workers in categories where wage differences are large enjoy a higher productivity than firms with workers in categories with small wage differences. A flat wage structure yields low incentives to work hard since little is to be gained by the individual from being productive, while the contrary could be expected in a worker category where the wage structure is steep. The hypothesis we shall test is if productivity in a firm is higher the larger is the number of workers in categories where wage differences are large. To this end, we specify the following explanatory variable:

$$wwd_k = \sum_{j=1} \frac{a_{jk}}{\sum_j a_{jk}} CofV_j \quad (4)$$

where wwd_k is a variable measuring the weighted wage difference among the worker categories represented in each firm *k*. With coefficients of variation for our worker categories $CofV_j$ (*j*=1…114) we can define for each firm a coefficient of variation weighted by the number of workers in each category. We assume that the higher is wwd_k in firm *k* the higher will be the level of productivity in that firm. For instance, a firm with a number of worker categories where wages differ a lot will get a high value of wwd_k, while a firm with a number of worker categories where wages do not differ much will get a low value of wwd_k. If incentives to work hard and be more productive are higher in the first group, we should expect a firm with a large number of such worker categories to be more productive.

Starting with a production function for a single firm we have the inputs of labor in efficiency units, $e^a L^b$ and capital, *K*, and to which we add a firm specific factor *f*:[16]

$$Q = (e^a L)^d K^{1-d} f \varepsilon \, . \quad (5)$$

ε is an error term uncorrelated with the inputs. To get at labor productivity, *P*, we divide both sides by total inputs of labor, *L*, to get

$$P = Q / L = (e^a L^b)^d K^{1-d} L^{-1} f \varepsilon \, . \quad (6)$$

Effort is assumed to be a function of the relative wage, wwd_k, and is specified as $e = wwd_k^{\alpha}$. In the empirical specification we differentiate labor into high skilled and low skilled to get the estimable function

$$\ln \mathrm{P}_{kt} = \alpha_0 + \gamma \ln wwd_{kt} + \beta Z_{ks} + \delta \ln L_{hkt} + \tau \ln L_{lkt} + \kappa \ln K_{kt} - \eta \ln L_{kt} + \varepsilon_t \, . \quad (7)$$

where $\gamma = \alpha a$ and *k* is an index of firms and *t* represent time. Z_{ks} is a dummy variable of firm *k* aimed at capturing specific properties of the sector to which firm *k* belongs. This would capture the major differences between for instance firms in service sectors and firms in different manufacturing sectors. The capital stock, K_k, is measured as the value of machinery and

[16] Cf. Levine (1992).

equipment, in firm *k*. Both these variables are included to control for the major determinants of productivity differences across firms, namely the factor supplies. Productivity is measured as the firm's value added, divided by the total number of workers in the firm.

In *Table 3* we present six different specifications. *Model 1* shows the results of an OLS regression with all variables measured in the same year, 2002. The estimated elasticity of the weighted wage difference variable is .23 and is highly significant suggesting that wage differences in a worker category raise productivity significantly in a firm using workers in that category. Also the estimates of the parameters of the two factor supply variables come out highly significant.

Could it be argued that firms' productivity can affect the wage differences within a given worker category? If this is the case, wwd_k would be affected by P_k and hence not be truly exogenous and our estimates would be biased. This would occur only if a productivity increase raises the *dispersion* of productivity that in turn would affect wage dispersion. Since we estimate the *level* of firms' productivity as a function of the worker category weighted wage distribution, such an effect seems far from likely. Yet, should a problem in this respect exist, a remedy for such a source of estimation bias could be to estimate (7) using $wwd_{k,t-1}$ rather than $wwd_{k,t}$. It could also be the case that the productivity effect emerges with a lag of one year. The results when the firms' productivity of the year 2002 is estimated as a function of weighted wage differences in 2001 are presented as *Model 2*. The estimated parameter comes out somewhat lower than for $wwd_{k,t}$, (.14), but is again highly significant and supports the general conclusion that productivity is favourably affected by wage dispersion.

An alternative specification involves the use of instrumental variables. Wage dispersion is, however, a variable not easily instrumented. The only possible formulation is to instrument $wwd_{k,t}$ using $wwd_{k,t-1}$. The correlation between these two variables is .75. In *Model 3,* we present the results of this IV-estimation. We see that the estimated parameter now is somewhat higher, .34. The estimate is again highly significant, supporting the hypothesis that the increased wage differences within many worker categories have affected productivity in a favourable way.

Table 3. Estimating the effects of wage dispersion within worker categories on firms' productivity. Estimated on double logarithmic form.

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant	5.9734 (23.17)	6.2971 (61.47)	7.8788 (61.29)	6.5212 (60.55)	6.0150 (22.32)	6.4459 (61.67)
$\mathbf{wwd_{kt}}$	.2294 (9.82)		.3417 (7.29)	.2077 (8.26)	.1828 (6.20)	.1472 (5.84)
$\mathbf{wwd_{kt-1}}$		.1422 (6.96)				
$\mathbf{wwdf_{kt}}$				.0303 (3.77)		.0147 (1.69)
$\mathbf{wd_{kt}}$					.0484 (3.07)	.0587 (4.60)
LS	.0792 (3.35)	.0447 (1.55)	.0844 (2.94)	.0664 (2.44)	.0750 (2.93)	.0573 (2.09)
HS	.0818 (7.38)	.0942 (7.15)	.0688 (5.10)	.0778 (6.66)	.0805 (7.06)	.0764 (6.57)
K	.1384 (17.34)	.1423 (15.69)	.1433 (15.90)	.1338 (15.61)	.1371 (16.48)	.1317 (15.32)
L	-.2616 (-8.75)	-.2616 (-7.02)	-.2629 (-7.20)	-.2577 (-7.71)	-.2737 (-8.71)	-.2561 (-7.64)
Sector dummies	Yes	Yes	Yes	Yes	Yes	Yes
No of obs.	3882	2901	2901	3348	3647	3348
$\mathbf{R^2}$	.3538	.3839	.3836	.3717	.3676	.3759

Notes: Robust estimates. The correlation between $wwd_{k,t}$ and $wwd_{k,t-1}$ is .75. 52 sector dummy variables have been used.

How large is this estimated effect of wage dispersion on productivity? The two estimations based on $wwd_{k,t}$ yield elasticities of .23 and .34, respectively. Labour productivity in all sectors increased by approximately 3 % per year between 1996 and 2002 while wage dispersion within categories increased by approximately 6.6 % per year. Our result in *Model 1* Table 3 would then suggest that 1.5 out of the 3 percent of the productivity gains should be ascribed increasing wage differences. This amounts to approximately half of the productivity increase. However, some caution should be taken in interpreting the results for such large changes in the wage dispersion.

To test if other distributional variables affect productivity of the firm and to see if the weighted wage distribution within worker categories yields a robust estimate, we include in Table 3 the results from estimating three other models. We may first ask if the wage distribution of a worker category *within the firm* matters to the firm's productivity. We may test if the fact that the wage distribution within the worker category affects productivity is because it reflects the distribution of the worker category within the firm. Instead of defining wage differences for the worker categories we thus do it for the worker categories in each firm by calculating:

$$wwdf_k = \sum_{j=1} \frac{a_{jk}}{\sum_j a_{jk}} CofV_{jk} \tag{8}$$

which differs from equation (4) only in that the coefficient of variation now has an index *k* attached to indicate that it is the wage distribution of the worker category in firm *k*. The coefficient of variation for a worker category within a firm can only be calculated for some minimum number of workers in that category and we restrict this number to five.

Model 4 shows the results when the variable specified in (8) is added and we see that the estimated parameter of the worker category wage distribution remains fairly stable and changes from a value of .23 (Model 1) to .21. The added variable, $wwdf$, comes out with the expected positive sign and is significant.

Does the distribution of wages *across worker categories in the firm* matter? In *Model 5* we have included a variable of the wage distribution in the firm, wd_k, which measures the coefficient of variation *across all employees* in the firm. In line with Akerlof and Yellen one could argue that a wider distribution of wages in the firm would have a negative impact on productivity. One could argue that a large wage distribution also across worker categories would stimulate productivity to the extent that workers can switch between worker categories. More importantly is that this effect may capture the fact that many positions, particularly at the upper tail of the wage distribution, may be held by persons from different worker categories. For instance, many management positions may be held by economists, legal advisers or engineers. Hence, the wage distribution across worker categories could matter to workers' incentives to perform well. One could also argue that some workers at the lower end would feel unfairly treated and lower productivity while those at the other end would be stimulated to work extra hard. The overall qualitative effect is therefore ambiguous. As seen in Table 3, *Model 5*, this variable yields a positive but low parameter estimate. We also see that our variable in focus, wwd_k, is robust also with respect to the inclusion of the distribution of all wages in the firm. The estimated parameter is now .18.

Finally, in *Model 6*, we include both the variable as specified in (8) and the variable on firm's wage distribution. We see that our estimated parameter of the variable representing the wage distribution across worker categories is slightly lower (.15) while the estimated parameter of wd_k is slightly higher. However, the variable $wwdf_{kt}$, measuring the effects of the distribution of wages within the worker categories in firms, yields an estimate that is still significant but now only on the 10 percent level.

Fair wage setting

The estimations presented above suggest that wage differences affect productivity. The theoretical justification for including the variables is simply that wage differences stimulate workers' effort. However, the wage-productivity relation should be more directly tested against efficiency wage theory. In general, inter-industry or inter-firm wage differentials have been shown to persist and to be unrelated to human capital differences or be due to compensating differentials.[17] The more explicit tests of efficiency wage setting are based on tests of the relation between relative wages and performance, including work satisfaction, absenteeism, quits etc. and most studies yield results consistent with efficiency wages.[18] Levine (1992) differs from other studies by presenting a direct test of a crucial implication of efficiency wage theory, i.e. if the increase in productivity is large enough to pay for the higher wage costs. He finds that the elasticity of output with respect to wages was of the magnitude predicted by efficiency-wage theories.

In this section we propose a slightly different test of the fair wage approach to efficiency wage setting. There are basically two major problems involved in the literature of empirical tests of fair wage models. One is that there is no definite notion of what constitutes the "fair" wage. Another, and more general problem, is that it may be hard to distinguish efficiency wage models from other models, notably the one that assumes that rent sharing affects wages. We argue that the available data and the suggested approach can handle these problems in an acceptable manner.

When testing for efficiency wages, Levine (1992) uses the relative wage of the firm, which in his data set is the average wage in the firm divided by the wage paid by its three largest competitors in the product market, controlling for occupation. Since we have access to linked worker-firm

[17] See, for instance, Krueger and Summers (1988), Groshen (1991) or Dickens and Katz (1987).

[18] Levine (1991), Akerlof et al (1988), Holzer (1989), Leonard (1987).

data, the situation is quite different. It is assumed here that each individual worker in the firm forms a norm of her fair wage from three different sources that all seem acceptable. The first source of fairness is the worker's *own individual characteristics*, like education, seniority at the firm, age etc. Perfectly in line with fair wage theorizing, a highly educated individual would feel unfairly treated if she were to be paid like an unskilled worker. Increasing seniority is also likely to add to the individual's perception of the fair wage. Age does not necessarily add to the perceived fair wage. While experience increases in age, aging could also be argued to lower productivity and hence not necessarily add to the perceived fair wage.

Gender is the most controversial argument and several alternatives are plausible and will be tested. Assume that there is a discriminatory gender effect. This could add to the norm and workers would not think of it as unfair if workers of either sex, *ceteris paribus*, received a lower wage than workers of the other sex. If both men and women think of a gender wage differential as fair, gender should then be a determinant of the fair wage that we shall predict.

On the other hand, if discrimination is perceived of as truly unfair, then gender should *not* be included in the determinants of the relevant variable. Both alternatives will be tested. It is also plausible that the better paid gender, i.e. males, perceive a higher wage for males to be part of the fair wage while the worse paid gender, females, do not.

The second determinant of fair wage concerns sector. Wage unfairness would be perceived if an individual would be paid less than the standard wage in the sector. The idea is that a norm for the sector is established and that workers expect high wages in some sectors and low wages in others. This implies that we need to account for sector specific effects.

The third argument to affect the wage norm is the firm's profitability. The individual would perceive the wage to be unfair if the rents of the firm were not reflected in her wage. This implies that rents would be incorporated into the perceived wage fairness. In this way, rent sharing becomes an integral part of efficiency wage setting.

The first argument involves standard individual characteristics of the Mincer equation. The first step in our approach is therefore to estimate a standard Mincer equation to which we then add the other arguments, i.e. sector and firm profitability. We then define the fair wage as the one *predicted* by an estimated Mincer equation. Thus, we would get the predicted fair wage w_i^f of individual *i* as

$$w_i^f = \tilde{\alpha}_0 + \tilde{\alpha}_1 Z_i + \tilde{\alpha}_2 S_i + \tilde{\alpha}_3 (R/K)_i + \varepsilon_t \qquad (9)$$

where the $\tilde{\alpha}$:s represents an estimated parameters of the respective variables, Z_i is the vector of individual characteristics of individual *i*, S_i is sector in which individual *i* is active, and $(R/K)_i$ is the firm's result per value of capital in the firm where the individual works.

According to fair wage version of efficiency wage theory, the relevant variable to the individual is w_i/w_i^f, i.e. the actual wage divided by the perceived fair wage. The hypothesis that we want to test is if this variable, in line with fair wage theory, affects productivity. To estimate the effects of this relative wage on productivity, we weight this variable for each firm *k* to get:

$$fw_k = \frac{1}{n_k}\sum_{j=1}^{n_k}(w_j / w_j^f) \qquad (10)$$

where n_k is the number of workers in firm *k*. (10) is our weighted fair wage variable. Hence, if a firm has a large number of workers with high wages compared to their "fair" wage, we expect this firm to show a high productivity. To test for this, we estimate

$$\ln \mathrm{P}_{kt} = \alpha_0 + \gamma \ln fw_{kt} + \beta S_{ks} + \delta \ln L_{hkt} + \tau \ln L_{lkt} + \kappa \ln K_k - \eta \ln L_{kt} + \varepsilon_t \qquad (11)$$

γ is the crucial parameter of interest. However, before proceeding to estimation, we need to be convinced that productivity *P* does not affect our right hand side variable *fw*. One could argue, in line with the rent sharing hypothesis, that productivity, our dependent variable, could affect the results per capital invested in firms and hence our explanatory variable *fw*. Given the determination of this variable as in (10), this does not seem likely. Nevertheless, we test the correlation between P_k and $(R/K)_k$ and find that this is not significant.[19] We therefore proceed towards estimating (11).

Table 4 shows the results of estimating (11) for the years 1999 and 2002. We select these years since 1999 was a year when wage differences increased a great deal, suggesting that this could be because firms' wage setters expected this to give rise to large productivity effects, while 2002 was a year when wage differences did not increase. This could suggest major differences in the estimated parameters γ of the two years.

In *Model 1*, the fair wage is based on an estimation of the Mincer equation using education, seniority, age and gender as the determinants of the fair wage. Including gender in the estimation of the Mincer equation that we use for predicting the fair wage, implies that a gender wage gap is considered fair by the individual workers. We see that for both years, an

[19] The coefficient of correlation is .0018.

increase in the individual's wage relative the fair wage has a positive effect on the firm's productivity. The effect differs a great deal and is considerably stronger, .90, in 1999 than in 2002, .45. The effect is thus only half as strong in 2002. The fact that firms and the individual workers determined the wages and that there were major increases in wage differences in the late 1990:s but not in the early 2000:s is consistent with the fact that the estimates are higher in 1999. Thus, the results suggest that there was a higher "pay-off" to an increase in the individual's wage relative the "fair" wage in this year than in 2002.

In general, alternative specifications of the "fair" wage could improve the fit of the model. That would imply better determinations of what represents the fair wage. As mentioned above, the gender variable is not at all an obvious determinant of the fair wage. A "feminist's" view would be that the gender wage gap should not be an integral part of the fair wage and we want to test if this is the case. *Model 2* is therefore based on an

Table 4. Estimating the effects of efficiency wage setting. Dependent variable is firms' productivity. Estimated on double logarithmic form. Robust estimates.

	Model 1 1999	**Model 1 2002**	**Model 2 1999**	**Model 2 2002**	**Model 3 1999**	**Model 3 2002**	**Model 4 1999**	**Model 4 2002**
Constant	5.8858 (80.05)	5.4462 (17.48)	5.8881 (80.18)	5.4341 (17.28)	5.9192 (80.29)	5.4512 (17.24)	6.3571 (59.58)	5.9228 (17.64)
fw_{kt}	.8970 (11.48)	.4500 (2.88)	.9133 (112.05)	.4595 (2.94)	.9282 (12.01)	.4666 (2.91)	.9105 (11.99)	.3934 (2.29)
wwd_{kt}							.1861 (5.78)	.1902 (7.72)
LS	-.0288 (-1.15)	.0460 (1.95)	-.0290 (-1.14)	.0464 (1.97)	-.0264 (-1.04)	.0477 (2.02)	.0015 (.06)	.0796 (3.31)
HS	.0906 (8.03)	.1081 (9.64)	.0937 (8.23)	.1104 (9.84)	.0904 (7.98)	.1085 (9.70)	.0662 (5.66)	.0857 (7.65)
K	.1495 (15.75)	.1325 (16.56)	.1465 (15.48)	.1311 (16.20)	.1476 (15,61)	.1316 (16.36)	.1477 (15.72)	.1334 (16.61)
L	-.2035 (-6.16)	-.2611 (-8.62)	-.2029 (-6.06)	-.2622 (-8.66)	-.2041 (-6.13)	-2645 (-8.66)	-.2040 (-5.96)	-.2678 (-8.91)
Sector dummies[2]	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No of obs.	3040	3882	3040	3882	3040	3882	3023	3882
R^2	.4159	.3537	.4182	.3547	.4185	.3551	.4328	.3679

estimation of the Mincer equation that *excludes* the gender variable as a determinant.[20] We note that the fit of the model now is slightly better in both years suggesting that the gender wage gap is *not* a part of the perceived fair wage. In general, the individual does not expect a gender wage gap to be fair.

This, however, is something that could differ across the two genders. Men, who generally are better paid than women, could very well be assumed to include a gender wage gap in their perception of the "fair" wage while women would not. In *Model 3*, we test for such a difference. We thus let the perceived fair wage of men include the estimated parameter times the value of the dummy variable for men, while for women the perceived fair wage includes the parameter multiplied by the average value of men and women. Would this specification further improve the fit of the model? We see that it does. The R^2 rises slightly, from .4182 to .4185 in 1999 and from .3547 to .3551 in 2002, when we allow gender to affect the determination of men's perception of the fair wage while women's perception of what constitutes a fair wage is assumed unaffected.

These results suggest that there are slight deviations in the perceptions of what constitutes a fair wage. It is encouraging that when the gender variable is neutralised, as in *Model 2*, this yields a better fit than when it is included since this suggests that it is not in general considered "fair" that gender should affect the wage rate. On the other hand, *Model 3* suggests that men still to some extent perceive gender to be a part of the determination of the fair wage.

Finally, in *Model 4*, we add the wage distribution variable analysed in Table 3 to see if this variable exerts an independent effect on productivity. We see that for both 1999 and 2002, a wider wage distribution as well as a larger wage/fair-wage ratio raises productivity. The added wage differentiation variable (*wwd*) is for 2002 now .19 while in the regressions in Table 3 it varied between .15 and .23 depending on specification. For 1999, the estimate is also .19. The estimate of the parameter of the wage/fair-wage variable remains relatively stable for 1999 (changes from .93 to .91) and for 2002 drops from .47 to .39.

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20 This means that we neutralise the gender effect. Assume $\hat{\gamma}$ to be the estimated parameter of the gender dummy variable in the Mincer equation. In determining the perceived fair wage we then include instead the value $\hat{\gamma}$ times the average value of the dummy variable for men and women (1.5).

Concluding remarks

The strong trend towards wage setting at the individual or local level among white-collar workers has been followed by increases in wage dispersion both within worker categories as well as across worker categories. Clearly, when local wage setters received more influence over wages, wage dispersion among white-collar workers increased and this pushed average white-collar workers' wages up. Wage differentials between white- and blue-collar workers increased. While these wage increases may have caused higher unemployment among white-collar workers, a favourable effect is the positive effects on labour productivity. Our results suggest that these productivity gains come about via more effort arising partly from prospects of better pay and partly from a higher wage relative the perceived fair wage, i.e. efficiency wage setting. The elasticity of productivity with respect to increases in the wage/fair-wage ratio seem to have diminished in 2002 as compared to 1999 which is as expected since the late 1990s was a period of drastically increasing wage differentials. Tests indicate that the relation between high relative wages and productivity is not due to rent sharing.

Twenty years of decentralisation of Swedish wage setting has implied that firms locally can affect the wage of the individual employee and firms' scope for efficiency wage setting has increased in a way previously not experienced.[21] Much of the changes towards decentralised wage setting has been motivated by employers' need for wage differences that should encourage the individual employee to invest in skills and to take own initiatives which is much in line with efficiency wage theory. It is conceivable that the extreme wage compression in Sweden up to the mid 1980s had adverse effects on work incentives and our results show that the wage differentiation, particularly in the late 1990s, triggered major effects on firms' productivity.

[21] Still, wage bargaining is a fact of life. It can easily be shown that a bargaining model with efficiency wage setting gives rise to higher unemployment than a bargaining model without efficiency wage setting or an efficiency wage model with no bargaining. See for instance Layard, Nickell and Jackman (1991), p. 540.

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Appendix 1

All data are taken from FIEFs data sets, which, in turn, derive from Statistics Sweden. To Statistics Sweden's annual wage investigations has been added wage data from "Kommunförbundet", Landstingsförbundet" and "Svenskt Näringsliv".

Wages cover the period 1995-2002 and includes a basic fixed wage (salary), any extra wage income like bonuses, any compensation for inconvenient work hours or compensation while "on duty", the value of fringe benefits, compensations in cash etc. All wages are expressed in month equivalents. White-collar workers cover the TCO/SACO areas and blue-collar workers the LO area.

Seniority is measured as number of years with the present employer.

Capital stock is measured as total value of machinery and inventories in the firm.

Share of highly educated is the number of employees with at least two years of college education as a share of all employees.

The total number of individuals is 2 261 514 and covers employees on permanent as well as temporary positions and includes entrepreneurs with employment conditions according to agreements. The number of positions is larger, 2 305 534, since some individuals have income from several jobs.

Appendix 2

Consider the Mincer model for year t

$$w_{ijt} = \gamma_t + a_t X_{ijt} + \beta_{t-1} WP_{jt-1} + \varepsilon_t \qquad \text{(A1)} \quad (4)$$

and for the year t-1

$$w_{ijt-1} = \gamma_{t-1} + a_{t-1} X_{ijt-1} + \beta_{t-2} WP_{jt-2} + \varepsilon_{t-1} \qquad \text{(A2)}$$

where we allow for changes in the parameters of the Mincer variables. Subtracting (A2) from (A1) we get

$$\Delta w_{ijt-1} = (w_{ijt} - w_{ijt-1}) = \alpha + a_t X_{ijt} - a_{t-1} X_{ijt-1} + \beta_{t-1} WP_{jt-1} - \beta_{t-2} WP_{jt-2} + \varepsilon\,, \qquad \text{(A3)}$$

where $\alpha = \gamma_t - \gamma_{t-1}$ and where t=2002 and t-1=1996. We include the gender and age variables (though gender is unchanged between the years and age differs additively) so as to allow for any parameter change during the six-year period. We lack data for seniority preceding 1996 as we do for changes in wage differences preceding 1996 (i.e. the variable WP_{jt-2}). The major changes in wage distribution took place after 1996. For this reason we instead include a constant α also in this specification. We include changes in education represented by dummy variables for individuals that increased their level of education during the period.

The results of estimating the model on this rate-of-change form are presented in *Table A1*. As is to be expected, the models can explain considerably less of the total variation in the change in wages as compared to the wage level (as in Table 2). Notable, however, is that when we add the wage differentiation variable based on the percentiles 90 and 10, as in *Model A2*, the adjusted R^2 rises considerably, from less than .05 to above .07. The explanatory power rises further, to above .08, when the coefficient of variation is included instead as in *Model A3*.

Table A1. Results for wage changes 1996-2002. Mincer wage equations. Dependent variable is the log of wage changes in month-equivalents. t-ratios. Edu$_{ij}$ represents individuals that have changed their education from level i to j during the period where 1=no high school, 2=2 years high school, 3=3 years high school, 4=less than 3 years university education, 5=at least 3 years university education, and 6= PhD.

	Model A1	**Model A2**	**Model A3**
Constant	8.5728 (601.41)	8.4620 (600.28)	8.5416 (609.50)
***Edu12**	-.3365 (-6.45)	-.2921 (-5.67)	-.2605 (-5.08)
Edu13	-.2042 (-13.48)	-.1651 (-11.04)	-.1502 (-10.09)
Edu14	.1269 (0.92)	.1239 (0.91)	.1293 (0.95)
Edu15	.2901 (2.30)	.3026 (2.42)	.2884 (2.32)
Edu23	-.2268 (-34.85)	-.2050 (-31.88)	-.1928 (-30.14)
Edu24	-.0579 (-1.94)	-.0676 (-2.29)	-.0710 (-2.42)
Edu25	.1300 (6.07)	.1192 (5.63)	.1077 (5.11)
Edu26	1.3336 (3.79)	1.4005 (4.03)	1.4485 (4.19)
Edu34	-.0381 (-4.89)	-.0545 (-7.08)	-.0570 (-7.45)
Edu35	.1100 (20.57)	.1042 (19.72)	.0889 (16.90)
Edu36	.6577 (7.93)	.5978 (7.30)	.6917 (8.49)
Edu56	.8462 (92.56)	.7888 (87.28)	.8841 (98.35)
Edu45	.4312 (73.07)	.4102 (70.35)	.4152 (71.55)
Edu46	.7154 (28.65)	.6144 (24.90)	.7297 (29.73)
Age	.01778 (28.31)	.0175 (28.22)	.0139 (22.53)
Age^2	-.0003 -(41.24)	-.0003 (-42.11)	-.0002 (-36.07)
Seniority	.0160 (11.52)	.0190 (13.84)	.0205 (15.05)

Seniority^2	-.0042 (-21.14)	-.0048 (-24.38)	-.0053 (-26.78)
Gender	-.2076 (-158.23)	-.1855 (-142.43)	-.2177 (-168.68)
$WP_j = PR_j$		2.3349 (176.42)	
$WP_j = CofV_j$			.5711 (209.87)
No of obs.	1268517	1268517	1268517
Adj R^2	.0483	.0711	.0803

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