Wage Linkages Between Private and Public Sectors in Sweden

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Abstract. The paper examines wage linkages between private and public sectors in Sweden by means of Granger causality tests and estimation of error correction models. Wage changes in central and local governments are Granger caused by private sector wage changes. Public sector wage increases involve error correction mechanisms; the lower the relative wage in the past, the higher the current wage increases. Increases in unemployment are associated with relative wage improvements for public sector employees.

1. Introduction

Recent theoretical and empirical work on wage determination has almost exclusively focused on wage setting in the private sector. This characterization appears to be particularly accurate for European labour market studies; to our knowledge, little attention has been paid to the driving forces behind public sector wage formation in European economies and to the wage linkages between private and public sectors. This neglect of the public sector is somewhat surprising considering that public sector employment has shown a marked upward trend in a number of countries.

Scandinavian thinking about intersectoral wage linkages has been heavily influenced by the "Scandinavian Model of Inflation", known in Sweden as the EFO-model.[1] A key assumption in this framework is

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that the tradeable sector (roughly corresponding to mining and manufacturing) acts as a wage leader with respect to the nontradeable sector. Nominal wage increases in the tradeable sector adjusts in the long run to the "room" for wage increases in the sector, the sum of productivity growth and international inflation. Wage increases in the tradeable sector are assumed to be transmitted to the rest of the economy through bargaining institutions and market forces.

Despite the popularity of the EFO-model, there have been few attempts actually to *test* the model's key wage leadership hypothesis.[2] Our paper offers evidence on this matter, focusing on wage linkages between private and public sectors in Sweden. We exploit quarterly data from the mid 1960s to the early 1990s, and examine in detail the pattern of sectoral wage movements.

We begin with a brief overview of some institutional facts of relevance. Section 3 proceeds to time series analyses involving Granger causality tests, cointegration regressions, and estimation of error correction models. The public sector is divided into central government and local governments. By and large, our results indicate that public sector wage increases are Granger caused by wage increases in the private sector.

Section 4 extends the previous analyses by introducing unemployment and consumer price changes as new variables in the regressions for public sector wage changes. Does unemployment matter for public sector wages? The answer according to our analyses is no; high unemployment is, therefore, associated with relative wage improvements for employees in the public sector as private sector wages tend to decrease when unemployment increases.

2. Background

The Swedish public sector has shown substantial growth over the past few decades. Employment in central and local governments amounted to 25 percent of total employment in 1965, whereas it stood at 41 percent of total employment in 1990 (Table 1). Public sector employment growth has been especially strong in the local government sector. During the second half of the 1980s, however, the local government employment share was constant.

The rise in public sector employment is accompanied by an increase in the share of public sector employees within the largest wage earner

organizations. Public sector employees represented 20 percent of the membership in LO – the largest organization – in 1960 and almost 40 percent in the late 1980s (Calmfors & Forslund, 1990).

Since 1965 employees in the public sector have had the same rights to negotiate over wages (including the right to strike) as private sector employees. Since then employer organizations for the central and local governments have regularly negotiated with public sector unions, with substantial coordination on both sides. There are three separate public sector employer organizations, one for central government and two for local governments (county councils and municipalities). Public sector employees are organized in three major organizations (LO, TCO, and SACO). The past few years have seen a move towards more decentralized wage setting within each bargaining area. Wage agreements in private and public sectors have usually been synchronized. With three exceptions, private sector negotiations have been completed before wage agreements in the public sector.

Public/private relative wages have displayed rather dramatic changes since the mid 1960s. Hourly wages – as computed from the National Accounts – have increased slightly faster in the private sector than in the public sector (Table 2). Average public sector wages were 20 percent higher than average private sector wages in the mid 1960s (Figure 1), a "wage premium" that was almost extinguished by 1976. Relative public sector pay increased until 1982 while the rest of the 1980s involved reduced public sector relative wages.

Table 1. Employment in private and public sectors, 1965-1990 (percentage of the total number of employees)

	Private sector	Central government	Local government
1965	75	12	13
1970	70	11	19
1975	66	11	23
1980	61	11	28
1985	58	11	31
1990	59	10	31

Note: The figures concern persons aged 16-74 (16-64 for 1990). Those employed in agriculture, forestry and fishing, as well as self-employed, have been excluded due to lack of data at the beginning of the period. In 1985 these groups amounted to nine percent of total employment.

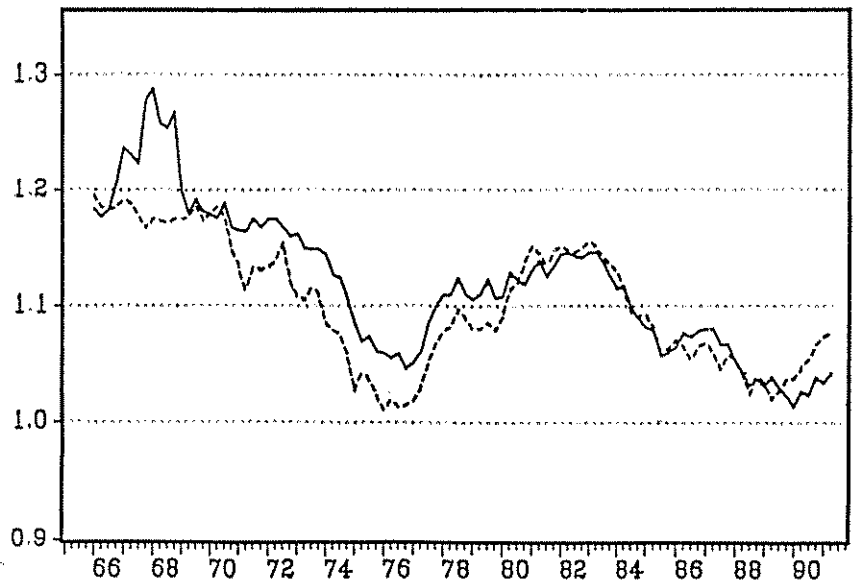
Source: The Swedish labour force surveys

Table 2. Annual nominal wage increases (percent) in private and public sectors, 1966-1990

	Private sector	Central government	Local government
1966-1973	8.3	7.5	8.2
	(1.1)	(1.5)	(4.8)
1974-1990	9.2	9.0	8.7
	(2.5)	(3.7)	(3.1)
1966-1990	8.9	8.5	8.5
	(2.2)	(3.3)	(3.7)

Note: Standard deviations within parentheses.
Source: See Appendix

Figure 1. The relative wages of the central government (dashed line) and the local governments (solid line) compared with the private sector, 1966:1-1990:2 (seasonally adjusted quarterly data)



Source: See Appendix.

3. Intersectoral Wage Linkages

We begin by examining the time series properties of the data. Table 3 presents the results of integration tests for the wage series as well as unemployment and consumer prices. The null hypotheses are that the series have a unit root, i.e. they are stationary when differenced. This is tested against the alternatives that the series are stationary or stationary around a deterministic trend (trend stationary).

The wage and price levels are obviously nonstationary. The ADF-tests with trend suggest that these series are difference stationary. The null hypothesis of a unit root cannot be rejected against the trend stationary alternative. For unemployment, on the other hand, the null hypothesis is rejected. We will therefore assume that the unemployment level is stationary in the subsequent analysis. The tests of the seasonal differences indicate that wage and price changes are stationary.[3]

We proceed to an examination of patterns of temporal precedence of wage changes in the different sectors making use of Granger causality tests. These tests can be described as follows.[4] Suppose we have two time series, $\{\Delta w_t^c\}$ and $\{\Delta w_t^p\}$. If we regress Δw_t^c on lagged values of Δw^c and Δw^p and find that the coefficients of the latter are zero then Δw^p fails to Granger cause Δw^c. The estimated model is:

$$\Delta w_t^c = \gamma_1 + \sum_{i=1}^{k} \alpha_i \Delta w_{t-1}^c + \sum_{i=1}^{k} \beta_i \Delta w_{t-i}^p + \varepsilon_t, \qquad [1]$$

where ε_t is a random error. An F-test may be used to test whether the β-coefficients are jointly zero. As an alternative, a t-test can be applied to test whether the sum of the β-coefficients is significantly different from zero.

The final prediction error suggested by Akaike (1969a and 1969b) offers a third possibility. The final prediction error is defined as:

$$\text{FPE} = \frac{n+q}{n-q} * \frac{1}{n} * \text{SSR}, \qquad [2]$$

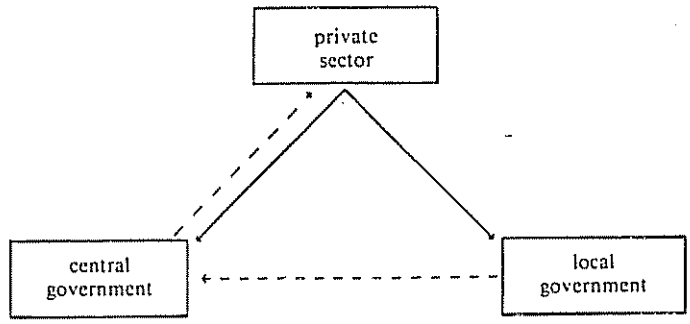
where n is the number of observations, q is the number of estimated parameters, and SSR is the sum of squared residuals. This is simply the estimated residual variance [SSR/(n-q)] multiplied by a penalty

Table 3. Integration tests, 1965:1-1991:2.
Significance levels are indicated by ** for the 1%-level and; * for the 5%-level

	SB, DW-test	DF, t-test Q(27), p-value		ADF, t-test Q(27), p-value	
		no trend	trend	no trend	trend
Levels					
Log wages in					
Central government	0.06	−0.23	−4.48	0.08	−1.45
		0.00	0.00	0.67	0.54
Local government	0.10	−0.46	−4.90	−0.90	−1.63
		0.00	0.00	0.28	0.22
Private sector	0.03	−0.49	−3.35	−0.36	−1.43
		0.00	0.00	0.54	0.53
Unemployment	0.43**	−2.61	−2.64	−3.25*	−3.79**
		0.00	0.00	0.99	0.93
Log consumer prices	0.00	1.79	−2.02	0.64	−2.84
		0.00	0.00	0.64	0.41
Seasonal differences					
Log wages in					
Central government	1.47**	−7.47**	−7.44**	−4.96**	−4.95**
		0.03	0.03	0.50	0.50
Local government	1.76**	−8.77**	−8.75**	−4.88**	−4.85**
		0.05	0.04	0.59	0.59
Private sector	1.07**	−5.92**	−5.84**	−4.21**	−4.20**
		0.08	0.07	0.42	0.42
Unemployment	0.36*	−2.81	−2.68	−4.19**	−4.08**
		0.10	0.10	0.53	0.53
Log consumer prices	0.16	−1.91	−2.05	−3.55*	−3.61*
		0.01	0.01	0.03	0.03

Notes: Seasonals are included. SB refers to Sargan & Bhargava (1983) where it is suggested that the Durbin-Watson statistic should be used to test for stationarity. Critical values are from Sargan & Bhargava (1983). DF and ADF refer to the Dickey-Fuller test and the augmented Dickey-Fuller test; see Dickey & Fuller (1979) and (1981). The t-values concern the null hypothesis that there is a unit root. The criteria for augmentation are white noise residuals and the significance of the lags. The number of lags are three in all cases except three, in these cases there are four lags. The p-values of the Ljung-Box Q-statistic for autocorrelation of the residuals, using 27 autocorrelations, are reported. The null hypothesis is that the residuals are white noise.

Figure 2. Causality of wage changes, 1966:1-1991:2



factor for additional parameters [(n+q)/n]. The final prediction error from the estimation of the (bivariate) equation [1] may be compared with that from a univariate specification (no β:s are estimated). If the inclusion of $\{\Delta w_t^p\}$ reduces the final prediction error we may conclude that $\{\Delta w_t^p\}$ Granger causes $\{\Delta w_t^c\}$. This procedure is similar to an F-test; our qualitative results are identical to those of the F-test.[5]

Table 4. Causality tests, 1966:1-1991:2

difference of log wages		F-test p-value	t-test p-value	FPE *100	FPE_a *100	n of lags	summary F-test	summary t-test
dep	indep						+ for causality	
central gov't.	private sector	0.009	0.072	0.266	0.286	4	+	+
local gov't.	private sector	0.050	0.010	0.138	0.142	10	+	+
local gov't.	central gov't.	0.351	0.712	0.155	0.141	14	–	–
central gov't.	local gov't.	0.082	0.978	0.290	0.292	11	+	–
private sector	local gov't.	0.227	0.848	0.081	0.077	9	–	–
private sector	central gov't.	0.011	0.110	0.071	0.077	9	+	–

Notes: The number of lags is determined by minimizing the final prediction error (FPE). The t-test concerns the sum of the estimated coefficients of the lags of the independent variable. FPE_a is the final prediction from an univariate estimation with the same lag length as for FPE.

The results from the estimations are reported in Table 4 and are summarized in Figure 2.[6] The two strong links go from the private sector to the public sectors. There are some indications of feedback between central government wage changes and private sector wage changes. No clear results emerge for the public sectors although the results suggest a link from local to central government.

The specification given by equation [1] can be seen as a special case, with restrictions on the parameters, of a more general specification including level variables. We now explore the consequences of relaxing these restrictions. In particular, we estimate error correction specifications of the form:

$$\Delta w_t^c = \gamma_1 + A[w_{t-4}^c - \lambda w_{t-4}^p] + \sum_{i=1}^{k} \alpha_i^{**} \Delta w_{t-i}^c + \sum_{i=1}^{k} \beta_i^{**} \Delta w_{t-i}^p + \varepsilon_t \qquad [3]$$

Engle & Granger (1987) propose a two step method for the estimation of specifications like [3]. The first step involves estimation of the long-run relationship:

$$w_t^c = \gamma_1^* + \lambda w_t^p + v_t, \qquad [4]$$

where v_t is a random error. If the series of estimated residuals, which is a linear combination of the wage levels, is stationary we may substitute it into [3] and estimate the following model as a second step:

Table 5. Cointegration regressions, 1965:1-1991:2
Significance levels for the stationarity test of the residuals are indicated by ** for the 1%-level and * for the 5%-level.

log wages		γ_1^*	λ	R^2	DW	DF, t-test Q(27), p-value	ADF, t-test Q(27), p-value n of lags
dep	indep						
central gov't.	private sector	0.214	0.956	0.994	0.940**	−5.778** 0.00	−1.658 0.45 3
local gov't.	private sector	0.331	0.934	0.995	1.396**	−7.443** 0.00	−3.424* 0.17 2
local gov't.	central gov't.	0.130	0.975	0.996	1.855**	−9.595** 0.00	−4.258* 0.75 3

Notes: Seasonals are included. Critical values for the Durbin-Watson statistic are from Sargan & Bhargava (1983). Critical values for the DF- and ADF-tests are from Engle & Granger (1987).

$$\Delta w_t^c = (\gamma_1 - \hat{\gamma}_1) + A\ \hat{v}_{t-4} + \sum_{i=1}^{k} \alpha_i^{**} \Delta w_{t-i}^c + \sum_{i=1}^{k} \beta_i^{**} \Delta w_{t-i}^p + \varepsilon_t \qquad [5]$$

where $\hat{\gamma}_1 = -\hat{\gamma}_1^* A$.

The results of the cointegration regressions are reported in Table 5. The DW-statistics and the DF-tests on the resulting linear combination of wages indicate that all regressions produce stationary combinations. The residuals of the DF-tests are, however, not white noise. Augmentation produces white noise residuals. The t-statistics of the ADF-tests are significant for the local government-private sector regression and the local government-central government regression suggesting that we have found two stationary linear combinations (or cointegrating relationships).

The ADF-test on the central government-private sector combination cannot reject the null hypothesis of a unit root. The test is appropriate for autoregressive time series of higher order. Said & Dickey (1984) discuss testing for unit roots for autoregressive-moving average time series (ARMA) models of unknown order. They suggest that a long sequence of autoregressive terms should be included in the test regression if one suspects that the series tested has nonzero MA-terms. If the ADF-test on the central government-private sector combination is augmented by 20 lags, the null of a unit root can be rejected at the 10 percent level, the t-value is –2.864. In addition, we have estimated an ARMA-model with one AR-term and two seasonal MA-terms. The estimated model gives white noise residuals (the p-value of the Q-test is 0.10) and the t-test for a unit root yields a value of –6.658. We therefore conclude that it is safe to proceed under the assumption that central government and private sector wages are cointegrated.

The results from the error correction specifications are reported in Table 6. Note that this specification (equation [5]) bears close resemblance to the specification of the Granger test (equation [1]). The only difference is that an error correction term (the deviations from the cointegrating relationships) is included. The problem addressed is thus whether the wage levels add anything to the structure found in the causality tests concerning wage changes. For comparison the results from the error correction estimations are reported together with the corresponding results from the causality tests previously reported in Table 4.

Table 6. Error correction regressions, 1966:1-1991:2

difference of log wages		$\Sigma\alpha_i^{**}$	$\Sigma\beta_i^{**}$	A	R^2	SEE	Q(27) p-value	n of lags
dep	indep							
central gov't.	private sector	0.222 (2.21)	−0.251 (1.40)	−0.652 (7.15)	0.41	0.044	0.28	1
		0.099 (0.55)	0.503 (1.82)		0.30	0.049	0.44	4
local gov't.	private sector	0.551 (1.93)	−0.403 (0.83)	−0.626 (3.66)	0.67	0.030	0.95	10
		0.001 (0.00)	0.935 (2.66)		0.60	0.033	0.62	10
local gov't.	central gov't.	0.318 (0.48)	0.283 (0.44)	0.041 (0.27)	0.62	0.034	0.68	14
		0.389 (0.55)	0.266 (0.37)		0.62	0.034	0.66	14
central gov't.	local gov't.	1.033 (1.70)	−0.494 (0.73)	−0.391 (2.31)	0.52	0.046	0.98	11
		0.614 (1.02)	0.018 (0.03)		0.48	0.048	0.98	11
private sector	local gov't.	0.331 (0.97)	0.109 (0.55)	0.159 (1.37)	0.52	0.025	0.90	9
		0.675 (2.89)	−0.033 (0.19)		0.51	0.025	0.89	9
private sector	central gov't.	0.419 (1.52)	−0.094 (0.62)	0.202 (2.66)	0.61	0.023	0.98	9
		0.928 (4.48)	−0.239 (1.62)		0.57	0.024	0.95	9

Notes: Absolute t-values in small types within parentheses. SEE refers to the standard error of the estimate. In the column headed Q(27) the marginal significance level of the Ljung-Box Q-statistic for autocorrelation of the residuals, using 27 autocorrelations, are reported. The null hypothesis is that residuals are white noise. Using a significance level of five percent, the null hypothesis is accepted for all estimations. N of lags refers to the number of lagged wage changes included, i e. the value of k in equation [5].

The error correction term has been lagged four quarters. The table reveals that the error correction formulations give "better" results than the causality formulations. The estimations where central and local government wage changes are regressed on private sector wage changes produce highly significant coefficients for the error correction variables. These coefficients indicate that more than 60 percent of deviations from the long run path are eliminated in one year. The error correction variable is also significant when central government wage changes are regressed on local government wage changes.

In conclusion, there is strong evidence that public sector wage changes are preceded (Granger caused) by private sector wage changes and that relative wage levels in the past influence current wage increases.

4. Extensions

We have so far focused exclusively on wage-interactions and ignored other variables that may influence wage increases in the public sector. We will conclude our investigation by examining the role of unemployment and price inflation. To what extent can public sector wage behavior be described by a simple Phillips-curve specification, augmented by an error correction term that captures past wage deviations relative to the private sector?

Table 7 displays the estimated equations for the central and the local governments. Price inflation is typically significant, whereas unemployment never enters significantly in any of the regressions. The error correction terms are always highly significant with estimated coefficients in a range between −0.4 and −0.6. The equations pass tests for nominal neutrality, i.e. wage inflation is homogeneous of degree one in the nominal variables.

The implications of these results for sectoral relative wage movements over the business cycle can be highlighted by a comparison with an estimated wage equation for the private sector. Many wage equations have been estimated for Sweden over the past few years, and they consistently suggest substantial wage responsiveness to unemployment. A representative example is the following equation, estimated by Holmlund (1988) on quarterly data for the Swedish private sector:

$$\Delta w_{p,t} = \underset{(6\,32)}{-0.455}\,[w_p - q]_{t-4} - \underset{(4\,00)}{0.056}\,u_{t-5} + \text{other variables}$$

The dependent variable is the four quarter change in the log real product wage (w_p), q is log labour productivity, and u is the log of the unemployment rate (t-ratios in parentheses). Increases in unemployment are followed by wage moderation in the private sector; in fact, several cross-country comparisons have indicated that Swedish wages are exceptionally responsive to unemployment.[7]

Table 7. Public sector wage equations

	central government		local government	
	unrestricted	restricted	unrestricted	restricted
excess relative wage	−0.436 (4.75)	−0.397 (4.45)	−0.607 (6.12)	−0.570 (5.70)
unemployment	−0.012 (1.39)	−0.007 (0.90)	−0.002 (0.29)	−0.003 (0.44)
lagged own wage changes	0.166 (1.77)	0.153 (1.62)	0.149 (1.82)	0.137 (1.64)
sum of changes in				
private sector wages	0.203 (1.07)	0.389 (2.58)	0.182 (1.30)	0.361 (3.11)
consumer prices	0.282 (1.63)	0.458 (3.39)	0.302 (2.12)	0.502 (4.45)
R^2	0.59	0.57	0.61	0.59
$\bar{R}^2$	0.54	0.53	0.56	0.54
SEE	0.038	0.039	0.030	0.030
Q(27), p-value	0.80	0.78	0.30	0.13
t-test, p-value	1.60, 0.11		2.22, 0.03	
H_0: sum of coefficients for nominal variables = 1				

Notes: The dependent variables are the four quarter changes in the public sector log wages. Absolute t-values within parentheses. Constant and seasonals included. The excess relative wage (the error correction term) and unemployment are lagged four quarters. The null hypothesis of white noise residuals cannot be rejected for any estimation.

Our basic finding, then, is the following. When labour market slack increases there is a fall in private sector wage changes but no direct effect on public sector wages. The short run effect is therefore an

increase in public sector wages relative to private sector wages. The resulting excess relative wage is gradually extinguished through the error correction mechanism.

It remains to be seen whether these patterns in the historical data will survive the institutional changes that have recently taken place in the wage bargaining system of the Swedish public sector. The increased emphasis on local wage determination may well make public sector wage behaviour more similar to wage behaviour in the private sector.[8]

5. Concluding Remarks

We have investigated intersectoral wage linkages in Sweden, focusing in particular on the links between private and public sectors. The main results are as follows:

1. Private sector wage changes Granger cause wage changes in central government as well as in local governments.

2. Public sector wage movements involve an error correction mechanism whereby past deviations from the long run relative wage path are offset; the lower the relative wage in the past, the higher the current wage increases.

3. High unemployment coincides with relative wage improvements for public employees compared to private sector employees.

4. The public sector wage equations exhibit nominal neutrality.

Our research thus finds little support for the view that wage increases are driven by wage push originating in the public sector. The results reveal a rather different pattern, where wage changes are transmitted from the private to the public sector. This is broadly consistent with the "Scandinavian Model of Inflation".

Appendix: the Data

Average hourly wage rates for each sector are computed from the National Accounts, using series on paid out wages and salaries and total hours worked among employees. The sources are the SNEP-Q data base (1965-1969) and Statistics Sweden (unpublished tables, 1970-1991).[9]

The unemployment rate is open unemployment according to the labour force surveys (Statistics Sweden), adjusted for 1987-1991 so that it conforms to the old definition (+0.5 percentage points).

The consumer price index is published by Statistics Sweden.

Notes

[1] The EFO model was developed by the research department directors of TCO (the Central Organisation of Salaried Employees), SAF (the Swedish Employers' Confederation) and LO (the Swedish Trade Union Confederation); see Edgren et al. (1973)

[2] A recent paper by Warginger (1991) is an exception.

[3] The variables studied are $w_t - w_{t-4}$, where w is the log wage, $p_t - p_{t-4}$, where p is the log consumer price, and $u_t - u_{t-4}$, where u is unemployment.

[4] The seminal paper is Granger (1969).

[5] We have also used the final prediction error to determine the lag length k. There exist several other criteria for determining lag lengths. In all cases considered here, Akaike's Information Criterion gives the same qualitative results as the FPE.

[6] We have added seasonals to model [1]. Excluding the seasonals does not change the qualitative results. In order to check the stability of our results, we have re-estimated the wage change equations for the period 1971-1988. The general picture does not change. Compared with Table 4 the only difference is some evidence of a link from local governments to the private sector.

[7] See for example Alogoskoufis & Manning (1988), Bean et al. (1986) and Layard et al. (1991).

[8] See Zetterberg (1990) for a detailed microeconometric analysis of public/private wage differentials. Zetterberg's study is confined to cross section data for 1974 and 1981 and does not illuminate movements in wage differentials between the 1960s and the 1970s.

[9] SNEP-Q is a quarterly econometric model of the Swedish economy developed at Uppsala University and FIEF (Trade Union Institute for Economic Research).

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