

Swedish Institute for Social Research (SOFI)

Stockholm University

WORKING PAPER 8/2004

IS THE SWEDISH CENTRAL GOVERNMENT A WAGE LEADER?

by

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November 22, 2004

Abstract

Is the Swedish central government a wage leader? This question is studied empirically in a vector error-correction model using a unique, high quality data set. Private sector salaries are found to be weakly exogenous to the system of equations. This means that the private sector is the wage leader in the long-run model. We also find that salaries in these two sectors do not converge to a common salary in the long-run and that changes in central government salaries do not Granger cause changes in private sector salaries. Together, these findings clearly demonstrate that the central government is not placing undue pressure on salaries in the private sector. The central government is not acting as a wage leader.

Keywords: public sector wages, Sweden, vector error-correction model, wage leadership.

JEL: C32, E24, J31.

*We would like to thank seminar participants at the Department of Economics, Stockholm University and the Swedish Institute for Social Research for their comments and suggestions. Matthew Lindquist is grateful to the Jan Wallander and Tom Hedelius Foundation for financial support.

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1 Introduction

The Scandinavian countries have small, open economies. Their labor forces are highly unionized and they tend to have relatively large public sectors. This particular combination of characteristics creates a unique set of challenges to the wage formation process. The main challenge is how to set wages in the highly unionized, non-competitive sectors without placing undue pressure on the wage formation process in the competitive sectors, pressure that would ultimately put these sectors at a significant disadvantage vis-a-vis their foreign competitors, particularly when exchange rates are fixed.¹

This problem has been widely recognized by politicians and trade union economists alike and was formalized in a number of economic models during the early 1970's. The Norwegian multi-sector price income model (Aukrust, 1970, 1977), the Swedish EFO-model (Edgren, Faxén and Odhner, 1973), and Finland's input-output framework (Halttunen and Molander, 1972) all address this problem explicitly. Collectively, these models are known as the Scandinavian model of inflation. The two main tenants of the Scandinavian model are; first, nominal wage changes in the competitive sector should be equal to the sum of productivity changes in the that sector plus changes in world prices and, second, that the competitive sector should act as the wage leader (i.e. wage increases should be transmitted from the competitive sector to the protected sector and not vice-versa).²

In Sweden, the EFO-model has been used by a number of economists to evaluate the wage formation process *ex post* (see e.g. Jacobson and Ohlsson, 1994 and Friberg, 2003). More importantly, it has acted as a set of normative guidelines for employers

¹If exchange rates are flexible, then upward pressure on wages in the competitive sector may result in currency depreciations. These automatic depreciations will increase exchange rate volatility. One could argue that there may be costs to doing business with a volatile exchange rate. Furthermore, total consumer welfare may go down by more than total producer welfare goes up when the exchange rate falls.

²Wage leadership can also be derived from institutional, wage bargaining models and efficiency wage models (see e.g. Bemmels and Zaidi, 1990).

and trade union negotiators, even after Sweden abandoned its fixed exchange rate regime. The normative conclusions of the EFO-model have been officially adopted by the Swedish Agency for Government Employers (Arbetsgivarverket) and guides their wage setting policies (Elvander, 2004; Lindquist and Vilhelmsson, 2004). The purpose of this paper is to examine whether or not actual wage outcomes of central government employees are in line with this stated praxis.³

We begin by presenting several institutional facts which may be relevant to the question at hand. First of all, central government wage agreements have, as a rule, been completed after wage agreements in the private sector have been signed (Holmlund and Ohlsson, 1992; Friberg, 2003; Elvander, 2004). Second, according to the Framework Appropriations System (Ramanslagssystemet) adopted in 1994, central government salaries are supposed to be explicitly tied to wage bill increases (net of average productivity growth) in the competitive sector. Third, the average salary of a central government worker is lower than that of a white-collar worker in the private sector (see Figure 1). Fourth, in 2002, the central government employed only 6 percent of all workers, while local government employed 28.5 percent and the private sector employed the remaining 65.5 percent.⁴ Together, these facts makes it less likely that the central government has been acting as a wage leader.⁵

A number of earlier studies concluded that the private sector was, in fact, the wage leader in Sweden (Holmlund and Ohlsson, 1992; Jacobson and Ohlsson, 1994; Andersson and Isaksson, 1997). This result is in line with the EFO-model and consistent with the stated goal of the Swedish Agency for Government Employers. How-

³It is not our goal in this paper to analyze the validity of the normative conclusions drawn from the EFO-model. These are taken as praxis. Our goal is to put this praxis to a rigorous statistical test.

⁴Central government employment peaked at 11.5 percent in 1979. Since then, it has been dropping steadily due to the privatization of a number of state owned companies, to the separation between church and state, and to the transfer of grade school and high school teachers from the central government to the local government. Source: Statistics Sweden (Statistiska Centralbyrån).

⁵See Lindquist and Vilhelmsson (2004) for a more thorough description of the relevant wage setting institutions as well as a short history of their development.

ever, several new reports published by the Swedish Central Bank (Tägtström, 2000; Friberg, 2003) have argued that the Swedish central government is now acting as a wage leader (at least for parts of the private sector). Our paper challenges these results and re-establishes the fact that there is no wage push coming from the central government.⁶

This is done using a unique, high quality data set, which is presented in Section 2. Unlike the previous studies by Holmlund and Ohlsson (1992), Jacobson and Ohlsson (1994), Tägtström (2000) and Friberg (2003), we do not compare the salaries of central government employees with the salaries of all workers in the private sector (i.e. an aggregate of blue-collar and white collar workers). Instead, we compare them to white collar salaries alone. We believe this to be the proper comparison, since more than 95 percent of all central government employees are white-collar workers (Lindquist and Vilhelmsson, 2004). Central government employees are members of white-collar unions and are, therefore, covered by white-collar contracts (Lindquist and Vilhelmsson, 2004). If the central government is putting undue pressure on the labor market, this effect should be most noticeable in the market for white-collar workers.⁷

The empirical results of this paper are presented in Section 3. They are based on the estimation of a vector error-correction model using the Johansen maximum likelihood approach (see e.g. Johansen, 1995). The methods used in this paper are similar to those used by Jacobson and Ohlsson (1994).⁸

We have three primary results. First, private sector salaries are found to be weakly

⁶Mizala and Romaguera (1995) test for public sector wage leadership in Chile. They find that after the deregulation of the Chilean labor market (between 1979-1982), the public sector lost its wage leading position.

⁷Andersson and Isaksson (1997) also make this distinction between white-collar and blue-collar workers, but their data set only goes up to 1995. Our updated data set allows us to consider the impact of the new Framework Appropriations System implemented in 1994 as well as the full impact of the move towards individual wage setting stipulated by the Framework Agreement (Ramavtal) which was put into place in 1990.

⁸Our impression is that Jacobson and Ohlsson (1994) were the first to apply the Johansen methodology in a stringent manner to construct a serious test of the EFO-model.

exogenous to the system of equations. This means that the private sector is the wage leader in the long-run model. Central government salaries adjust to changes in private sector salaries in order to maintain the long-run equilibrium relationship. Second, changes in central government salaries do not Granger cause changes in private sector salaries. Changes in private sector salaries are determined by a deterministic trend (domestic inflation) and a stochastic trend (which we interpret in line with the EFO-model as the sum of changes in exogenous productivity and changes in exogenous world prices). Third, we find that salaries in these two sectors do not converge to a common salary in the long-run. Together, these findings tell us that actual wage bargaining outcomes for central government workers are in line with the stated intentions of the Swedish Agency for Government Employers and that they are not placing undue pressure on the private sector market for white-collar workers.⁹

2 Data

We use two data series in this study: nominal monthly, white-collar salaries in the private sector, w_t^{ps} , and nominal monthly salaries in the central government, w_t^{cg} (see Figure 1). The data are annual time series from 1970 to 2002 collected by the

⁹We believe that the results presented in Tägtström (2000) and Friberg (2003) are the product of unfortunate choices of methods and data. Tägtström applies a standard Granger causality test to nonstationary data. These tests are (at best) only approximately correct and demand the use of non-standard F-test statistics. They may not be valid at all (Charemza and Deadman, 1992). Furthermore, when she tests for Granger causality using the data in first differences (i.e. using stationary data) these new, more correct tests show that the central government is not a wage leader.

Friberg (2003), on the other hand, uses methods similar to those employed in this paper and in Jacobson and Ohlsson (1994). The main methodological shortcomings of his paper is that he does not perform joint tests of cointegration and model specification, nor does he consider the impact of including deterministic components on the distribution of his test statistics. Given the large number of alternative models that he presents (and an equally large number that he fails to address), the fact that he does not test for model specification means that he does not reject a number of misspecified models nor does he necessarily find the most appropriate model. At the end of the day, our inability to distinguish between alternative models makes his results unambiguous and difficult to use in practice. The Granger causality tests in his paper are also incomplete. They examine the impact of x on z , but ignore the impact that x might have on z through a third variable y .

Confederation of Swedish Enterprise (Svenskt Näringsliv) and the Swedish Agency for Government Employers. They are based on actual contracts and cover nearly all workers in these two categories.

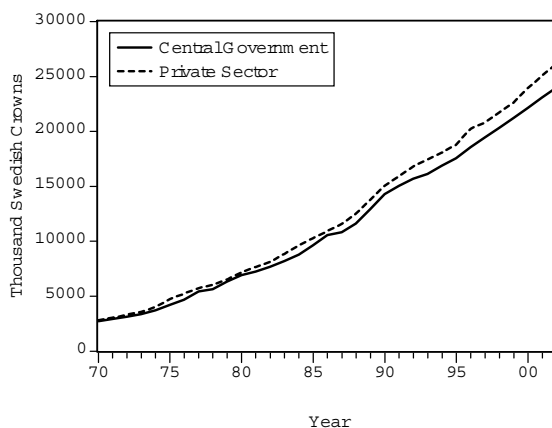


Figure 1: Salaries of White-Collar Workers, 1970-2002.

There are three major advantages of using this data, as compared to the data used in earlier studies. First, and most importantly, since more than 95 percent of all central government workers are white-collar workers, and since these workers are covered by white-collar unions, negotiations and contracts, it seems only reasonable to examine the impact of central government wage formation on wages of white-collar workers in the private sector. Comparing central government wages to an aggregate of white-collar and blue-collar workers in the private sector may be grossly misleading.

Second, the wage data used in this study have been correctly periodicized. For example, retroactive wages have been book-kept as yesterday's wages, whereas in the wage data from Statistics Sweden (Statistiska Centralbyrån) they are treated as today's wages. This type of periodicization is made possible by the fact that the data comes from employers with more precise knowledge about contracts and actual wages paid out.

Third, the two time series have been cleansed of between sector wage changes

due to structural changes. This is necessary because a number of large government companies have been privatized during this time period, including; the postal service, the telephone company, the largest energy producer, the railroad track maintenance company, and even the Swedish Lutheran church. Such changes in the underlying structure of the two sectors have been controlled for when producing the time series. Another important example is that primary and secondary school teachers are no longer central government employees. This has also been controlled for.

2.1 Pre-Testing the Data

Examining the time series for private sector, white-collar salaries and central government salaries in Figure 1, we see that both variables are clearly nonstationary. When this is the case, it is important to investigate the nature of this nonstationarity. To do this, we pre-test each variable in order to determine its order of integration (i.e. the presence of one or more unit roots) and to test for the presence of deterministic trends. This is done using the augmented Dickey-Fuller sequential procedure outlined in Enders (2004). The details of these tests can be seen in Appendix A.

The results of this sequential testing procedure are unambiguous. Both variables have a single unit root and are, hence, integrated of order one. Each of the variables also contains a quadratic deterministic trend which is due to the high level of inflation in Sweden during the 1970's and '80's. The fact that both variables are $I(1)$ means that they are potentially cointegrated. A joint test for cointegration and the presence of a quadratic trend in the preferred model will be carried out below. The results from this test tell us that we can, in fact, use regression analysis to say something meaningful about the relationship between these two variables despite the fact that they are both nonstationary and include stochastic trends.

3 Estimating a Vector Error-Correction Model

The empirical results of this paper are based on the estimation of a vector error-correction (VEC) model using the Johansen maximum likelihood approach (see e.g. Johansen, 1995). The VEC modelling strategy allows us to test for wage leadership in two distinct ways. First, we can examine if one of the variables included in the model is, in fact, weakly exogenous to the estimated system of equations. If two variables, X_t and Y_t , are cointegrated, and if the variable X_t turns out to be weakly exogenous, while the variable Y_t is not, then we know that the variable Y_t adjusts to changes in the variable X_t in order to maintain the long-run equilibrium. In this case, X_t is the "leader" and Y_t is the "follower". Second, the model allows us to construct a more robust test of Granger causality between ΔX_t and ΔY_t .¹⁰ One which does not suffer from the exclusion of a very important variable, namely, the long-run cointegrating relationship between X_t and Y_t .

We will also take advantage of the fact that a VEC model allows us to model both the short- and long-run relationship between white-collar salaries in these two sectors. This allows us to test for the presence of salary convergence in the long-run and to examine the determinants of salary formation in the short-run.

3.1 Determining the Lag Length

The first step in building a well specified, vector error-correction model is to determine the number of lags, p , which should be included in the model. This is done by estimating an unrestricted vector autoregressive model (VAR) model using the data

¹⁰ Δx denotes the first difference of variable x .

in levels.¹¹ The VAR(p) model can be written as

$$\mathbf{x}_t = \boldsymbol{\mu} + \mathbf{A}_1\mathbf{x}_{t-1} + \mathbf{A}_2\mathbf{x}_{t-2} + \dots + \mathbf{A}_p\mathbf{x}_{t-p} + \boldsymbol{\varepsilon}_t \quad (1)$$

where $\mathbf{x}_t = [w_t^{ps} \ w_t^{cg}]'$, $\boldsymbol{\mu}$ is a (2×1) vector with potentially nonzero constants. Each \mathbf{A} is a (2×2) matrix of regression coefficients and $\boldsymbol{\varepsilon}_t$ is a (2×1) vector of Gaussian, white noise error terms. This VAR(p) system of equations can be viewed as a model in reduced form. When determining the lag length p our goal is to obtain a parsimonious representation of the model which, at the same time, includes a sufficient number of lagged \mathbf{x}_t s so as to glean out all information available from the $\boldsymbol{\varepsilon}_t$ s concerning the structure of the relationship between the \mathbf{x}_t s. This means that our choice of p should be as minimal as possible, while, at the same time, we cannot allow non-normality, serial autocorrelation, or ARCH to appear in the residuals.

Following Enders (2004) we use multivariate generalizations of the Akaike information criterion (AIC) and Bayesian information criterion (BIC) to choose the appropriate lag length, p . The principle behind these two tests is the same. We are punished for adding variables that do not contribute significantly to the model fit. Oftentimes, these two test result in conflicting conclusion. But here they do not. Both the AIC and BIC choose $p = 1$ to be the appropriate lag length. This finding is confirmed by a set of likelihood ratio tests (Sims, 1980) which tell us that a VAR($p > 1$) model does not significantly outperform the VAR(1) model. The results

¹¹Since only lagged values of the endogenous variables appear on the right-hand side of each equation, there is no issue of simultaneity. Furthermore, the VAR model is "balanced". That is, the same regressors appear in each equation. Thus, OLS is an appropriate estimation technique. It can be applied to each equation in the system separately.

of these tests are reported in Table 1.

Table 1: Determining the lag length, p .

p	AIC	BIC	LR-test statistic	$\chi^2(4(p-1))$ -value (5%)
1	26.85	27.12		
2	26.93	27.39	2.43	9.49
3	27.20	27.85	2.64	15.51
4	27.37	28.21	3.63	21.03
5	27.73	28.78	3.22	26.30

Unfortunately, the residuals in the second equation of the VAR(1) model (i.e. the equation with central government salaries as the dependent variable) are not normally distributed. The Jarque-Bera test for normality has a p -value of 0.003. This implies that there is more information about the structure of the relationship in the data which we have not yet extracted from the residuals.

We continue by estimating a VAR(2) model. The residuals from this model are normally distributed, they do not suffer from serial autocorrelation nor do we detect the presence of ARCH. Thus, we accept $p = 2$ as the lag length in our VAR model. Table 2 shows that the AIC, BIC, and LR-tests all choose a VAR(2) model as the appropriate model given that we cannot accept a VAR(1) model due to non-normality of the residuals. With $p = 2$ in hand we can write down the unrestricted VAR(2) model as

$$\mathbf{x}_t = \boldsymbol{\mu} + \mathbf{A}_1\mathbf{x}_{t-1} + \mathbf{A}_2\mathbf{x}_{t-2} + \boldsymbol{\varepsilon}_t. \quad (2)$$

Table 2: Determination of lag-order, p .

p	AIC	BIC	LR-test statistic	$\chi^2(4(p-2))$ -value (5%)
2	26.93	27.39		
3	27.20	27.85	0.66	9.49
4	27.37	28.21	2.09	15.51
5	27.73	28.78	2.04	21.03

3.2 A Joint Test for Cointegration and Model Specification

The VAR(2) model can be rewritten in error-correction form as a VEC(1) model

$$\Delta \mathbf{x}_t = \boldsymbol{\mu} + \boldsymbol{\pi} \mathbf{x}_{t-1} + \boldsymbol{\pi}_1 \Delta \mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t. \quad (3)$$

Testing for cointegration between the nonstationary variables, \mathbf{x} , amounts to determining the rank of the matrix $\boldsymbol{\pi}$. If the rank of $\boldsymbol{\pi}$ is zero, then there are no linearly independent combinations of the nonstationary variables which are stationary. Thus, the nonstationary variables are not cointegrated. If the rank of $\boldsymbol{\pi}$ is two, then the variables themselves are both stationary (and the test for cointegration becomes redundant). If the rank of $\boldsymbol{\pi}$ is one, then there is one linearly independent combination of the nonstationary variables which is stationary. This means that the nonstationary variables are cointegrated. Thus, we want to test the hypothesis that $\text{rank}(\boldsymbol{\pi}) = 1$.

The Johansen method requires that we determine the rank of $\boldsymbol{\pi}$ and test for the presence of deterministic components in the model jointly, since the presence of deterministic components in the model affects the properties of the test for cointegration. To make these notions more clear, let us start by examining a more general version

of the VEC(1) model

$$\begin{aligned}
\Delta \mathbf{x}_t &= \boldsymbol{\mu}_{sr} + \boldsymbol{\delta}_{sr}t + \tilde{\boldsymbol{\pi}}\tilde{\mathbf{x}}_{t-1} + \boldsymbol{\pi}_1\Delta \mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \\
&= \boldsymbol{\mu}_{sr} + \boldsymbol{\delta}_{sr}t + \underbrace{\boldsymbol{\alpha}}_{\tilde{\boldsymbol{\pi}}} \underbrace{\begin{bmatrix} \beta_{ps} \\ \beta_{cg} \\ \mu_{lr} \\ \delta_{lr} \end{bmatrix}}_{\tilde{\mathbf{x}}_t}' \underbrace{\begin{bmatrix} w_t^{ps} \\ w_t^{cg} \\ 1 \\ t \end{bmatrix}}_{\tilde{\mathbf{x}}_t} + \boldsymbol{\pi}_1\Delta \mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t
\end{aligned} \tag{4}$$

where $\boldsymbol{\mu}_{sr}$ is a (2×1) vector of constants in the short-run model, $\boldsymbol{\delta}_{sr}$ is a (2×1) vector of regression coefficients which allow for a deterministic time trend, t , in the short-run model. The matrix $\boldsymbol{\pi}$ and the vector of variables \mathbf{x}_{t-1} are both modified to allow for the presence of a single constant, μ_{lr} , and a single deterministic time trend, $\delta_{lr}t$, in the long-run model (i.e. in the cointegrating vector). These are denoted as $\tilde{\boldsymbol{\pi}}$ and $\tilde{\mathbf{x}}_{t-1}$, respectively. The matrix $\tilde{\boldsymbol{\pi}}$ can be factored into a (2×1) vector, $\boldsymbol{\alpha}$, which represents the speed of adjustment to the long-run equilibrium and a (1×4) vector $\boldsymbol{\beta} = [\beta_{ps} \ \beta_{cg} \ \mu_{lr} \ \delta_{lr}]$ that represents the long-run (equilibrium) cointegrating vector. This general VEC(1) model encompasses 5 distinct models:

model 1: $H_0 : \boldsymbol{\mu}_{sr} = \mu_{lr} = \boldsymbol{\delta}_{sr}t = \delta_{lr}t = 0$

model 2: $H_0 : \boldsymbol{\mu}_{sr} = \boldsymbol{\delta}_{sr}t = \delta_{lr}t = 0$

model 3: $H_0 : \mu_{lr} = \boldsymbol{\delta}_{sr}t = \delta_{lr}t = 0$

model 4: $H_0 : \boldsymbol{\delta}_{sr}t = 0$

model 5: H_0 : no restrictions on the deterministic components.

Our task is to identify which of these models fits the data best at the same time as we test for cointegration. We can do this by testing different sets of restrictions jointly with the restriction that the rank of $\tilde{\boldsymbol{\pi}}$ is either 0, 1, or 2. We can minimize on

the number of tests necessary to complete this task by realizing that neither Model 1 or Model 2 are reasonable representation of the data, since the data trends upwards over time. This trend can be captured in model 3 by allowing for a non-zero drift term in each equation, $\boldsymbol{\mu}_{sr}$. Models 4 and 5 are also reasonable representations of the data. Model 5, however, is the only model which explicitly allows for a quadratic, deterministic trend in the data, which is what we found when we pre-tested the variables. We can also exclude the test for $\text{rank}(\tilde{\boldsymbol{\pi}}) = 2$, since both variables are $I(1)$.

This leaves us with a set of 6 joint null hypotheses to be tested. These null hypotheses can be ordered from the most restrictive test to the least restrictive test as follows: model 3 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 0$; model 4 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 0$; model 5 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 0$; model 3 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 1$; model 4 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 1$; model 5 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 1$. Table 3 shows each of these null hypotheses along with the appropriate likelihood-ratio (trace) test.

Table 3: Joint Determination of $\text{Rank}(\tilde{\boldsymbol{\pi}})$ and Deterministic Components.

$H_0 :$	Model 3		Model 4		Model 5	
$\text{rank}(\tilde{\boldsymbol{\pi}})$	LR-stat	5%	LR-stat	5%	LR-stat	5%
0	19.44*	15.41	27.84*	25.32	18.81*	18.17
1	5.46*	3.76	10.49	12.25	1.47	3.74

* denotes rejection of H_0 at 5% significance level.

Critical values taken from Osterwald-Lenum (1992).

Four of the six null hypotheses are rejected at the 5 percent significance level. Although H_0 : model 4 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 1$ is not rejected by the likelihood-ratio test, the residuals from this model trend upwards. As we saw in the pre-tests of the variables, there is a deterministic trend in $\Delta \mathbf{x}_t$. Model 4 restricts this trend to be zero and, hence, the trend in $\Delta \mathbf{x}_t$ shows up in the residuals.

H_0 : model 5 \cap $\text{rank}(\tilde{\boldsymbol{\pi}}) = 1$ is not rejected. The residuals from this model are

normally distributed, not serial autocorrelated, nor do they suffer from ARCH. The AIC and BIC also choose model 5 over model 4. Thus, model 5 is our preferred model.

The rank of the estimated $\tilde{\pi}$ matrix, $\hat{\pi}$, is equal to one which means that the long-run model $\tilde{\pi}\tilde{\mathbf{x}}_{t-1}$ is indeed cointegrated. The dual of this result is that there is one common stochastic trend driving the long run model. This stochastic trend is often assumed to be the sum of exogenous domestic productivity and exogenous world market prices (see e.g. Jacobson and Ohlsson, 1994). The quadratic, deterministic trend can be interpreted as domestic inflation, where the quadratic part is due to the high level of inflation in Sweden during the 1970's and '80's.

3.3 Testing Structural Hypotheses

In this section of the paper, we are interested in testing two hypotheses. First, and most importantly, does the central government act as a wage leader? Second, do white-collar salaries in different sectors converge over time to a common salary?

These two structural hypotheses can be formulated as restrictions on the VEC(1) model and then tested. To do this, we factor the (2×4) matrix $\hat{\pi}$ into a (2×1) vector, $\hat{\alpha}$, and a (4×1) vector, $\hat{\beta}$, such that $\hat{\pi} = \hat{\alpha}\hat{\beta}'$. The first vector, $\hat{\alpha} = [\hat{\alpha}_{ps} \ \hat{\alpha}_{cg}]$, which is often referred to as the "loading" matrix, is a pair of weights concerning the importance of the long run relationship (cointegrating vector) in explaining changes in $\mathbf{x}_t = [w_t^{ps} \ w_t^{cg}]'$. The coefficients in $\hat{\alpha}$ measure the speed of adjustment to past equilibrium errors. The second vector, $\hat{\beta}' = [\hat{\beta}_{ps} \ \hat{\beta}_{cg} \ \hat{\mu}_{lr} \ \hat{\delta}_{lr}]$, is the cointegrating vector itself, which defines the long run equilibrium relationship between w_t^{ps} and w_t^{cg} .

The estimated values of $\hat{\alpha}$ and $\hat{\beta}$ are reported in Table 4. The values of $\hat{\beta}$ have been normalized with $\hat{\beta}_{ps}$. The standard errors of those coefficients in $\hat{\beta}$ that are not uniquely identified are not reported. Testing structural hypotheses amounts to

testing restrictions on $\hat{\alpha}$ and $\hat{\beta}$.¹²

Table 4: Estimates of $\hat{\alpha}$ and $\hat{\beta}$.

$\hat{\beta}_{ps}$	$\hat{\beta}_{cg}$	$\hat{\mu}_{lr}$	$\hat{\delta}_{lr}$	$\hat{\alpha}_{ps}$	$\hat{\alpha}_{cg}$
1	-1.17	50.6	66.3	-0.013	0.769
	(0.039) ^a			(0.292)	(0.260)

a) Standard errors in parentheses.

3.3.1 Testing for Wage Leadership in the Long-Run Model

Testing for wage leadership in the long-run model amounts to testing each variable for weak exogeneity. The existence of only one cointegrating vector simplifies this test: we need only examine the t -values associated with $\hat{\alpha}_{ps}$ and $\hat{\alpha}_{cg}$. These are -0.044 and 2.96, respectively. Since we cannot reject $H_0 : \alpha_{ps} = 0$, we conclude that w_t^{ps} is weakly exogenous to the system of equations. Central government salaries, w_t^{cg} , on the other hand, are endogenous to the system of equations, since we can clearly reject $H_0 : \alpha_{cg} = 0$.

The null hypothesis that the central government acts as a wage leader ($H_0 : \hat{\alpha}_{ps} \neq 0 \cap \hat{\alpha}_{cg} = 0$) is strongly rejected. On the other hand, we cannot reject the hypothesis that the private sector is the wage leader ($H_0 : \hat{\alpha}_{ps} = 0 \cap \hat{\alpha}_{cg} \neq 0$). The test for weak exogeneity shows us that adjustments to the long run equilibrium are made through adjustments to central government salaries. That is, central government salaries react to changes in private sector salaries. They (alone) uphold the long-run relationship between the two sectors. In fact, $\hat{\alpha}_{cg} = 0.769$ tells us that the central government corrects 77 percent of the equilibrium error within one year's time. Together, these tests tell us unambiguously that the private sector is the wage leader and the central government is the wage follower.

¹²As we shall see, this can be done without first identifying (uniquely) all of the coefficients in $\hat{\beta}$.

3.3.2 Testing for Wage Leadership in the Short-Run Model

Given that w_t^{ps} is weakly exogenous to the system of equations, we can factorize the model into two single equations: one marginal model of Δw_t^{ps} and one conditional model for Δw_t^{cg} .¹³ We can use the marginal model of Δw_t^{ps} to test whether or not Δw_t^{cg} Granger causes Δw_t^{ps} . This can be viewed as a test for wage leadership in the short-run model.

Estimating the marginal model of Δw_t^{ps} results in the following regression equation

$$\Delta w_t^{ps} = \frac{267}{(103.5)} + \frac{20.5}{(6.771)} t + \frac{0.02}{(0.245)} \Delta w_{t-1}^{ps} + \frac{0.17}{(0.233)} \Delta w_{t-1}^{cg} + \varepsilon_t^{ps}. \quad (5)$$

An F -test concerning the hypothesis that the two coefficients in π_1 are equal to zero has a p -value of 0.58. So, we can pare down the marginal model to

$$\Delta w_t^{ps} = \frac{267}{(103.5)} + \frac{20.5}{(6.771)} t + \varepsilon_t^{ps} \quad (6)$$

which has an $\overline{R}^2 = 0.54$. The residuals are normally distributed and do not suffer from serial autocorrelation or ARCH. Thus we can conclude that Δw_t^{cg} does not Granger cause Δw_t^{ps} . Changes in private sector salaries are determined by a deterministic trend (domestic inflation) and by a stochastic trend (domestic productivity + world market prices) and not by changes in central government salaries.

3.3.3 Testing for Wage Equalization

Testing for long-run convergence in salary levels between sectors amounts to a test for homogeneity. A test for salary convergence can be formulated as, $H_0 : \widehat{\beta}_{ps} = 1 \cap \widehat{\beta}_{cg} = -1$ (where both coefficients are first normalized by $\widehat{\beta}_{ps}$). The alternative hypothesis is formulated as; $H_A : \widehat{\beta}_{cg} \neq -1$.

Since w_t^{ps} is weakly exogenous to the system of equations, we can carry out this

¹³This also means that we can estimate each equation separately using OLS.

test by estimating a single equation for Δw_t^{cg} . Estimating the conditional model of Δw_t^{cg} results in the following regression equation

$$\Delta w_t^{cg} = \underset{(97.18)}{254} + \underset{(32.41)}{87.1} t + \underset{(0.288)}{0.67} w_{t-1}^{ps} - \underset{(0.320)}{0.83} w_{t-1}^{cg} - \underset{(0.275)}{0.11} \Delta w_{t-1}^{ps} + \underset{(0.268)}{0.42} \Delta w_{t-1}^{cg} + \varepsilon_t^{cg}. \quad (7)$$

The t -value for $\pi_{1,1}$ equals 0.40. So, we can pare down the conditional model of Δw_t^{cg} to

$$\Delta w_t^{cg} = \underset{(90.29)}{241} + \underset{(34.41)}{84.8} t + \underset{(0.288)}{0.60} w_{t-1}^{ps} - \underset{(0.256)}{0.75} w_{t-1}^{cg} + \underset{(0.169)}{0.33} \Delta w_{t-1}^{cg} + \varepsilon_t^{cg} \quad (8)$$

which has an $\bar{R}^2 = 0.55$. The residuals are normally distributed and do not suffer from serial autocorrelation or ARCH. The normalized, long-run relationship (cointegrating vector) is given by $[1 \ -1.25 \ 241 \ 84.8]$.¹⁴ The speed of adjustment parameter, $\hat{\alpha}$, is now equal to 0.75.

The null hypothesis of the homogeneity test is $H_0 : \hat{\beta}_{cg}/\hat{\beta}_{ps} = -1$. The Wald test statistic for this restriction has a p -value of 0.00. We must, therefore, reject the null hypothesis of wage homogeneity. Salaries in the two sectors do not converge over time. This can also be seen in figure 2 which shows a simple plot of the ratio of private sector to central government salaries, w_t^{ps}/w_t^{cg} .¹⁵

4 Conclusions

This paper clearly shows that there is no wage push coming from the central government. The central government is not acting as a wage leader. This result reaffirms previous findings by Holmlund and Ohlsson (1992) and Jacobson and Ohlsson (1994), but stands in stark contrast to two recent papers published by the Swedish Central

¹⁴The estimates of the constant and the linear trend are actually conglomerate estimates of the constants and trends in both the short-run and the long-run conditional model.

¹⁵It is important to keep in mind, however, that the data has in no way been cleansed of potential changes in the composition and characteristics of the two different groups of workers. So, w_t^{ps}/w_t^{cg} can not be interpreted as a standard wage premium.

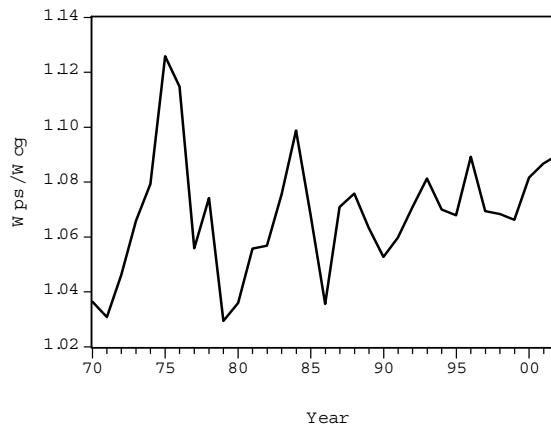


Figure 2: Ratio of Privates Sector Salaries to Central Government salaries.

Bank (Tägtström, 2000 and Friberg, 2003). This paper also finds that the salaries of white-collar workers in the private sector and central government show no tendency to converge to a common salary in the long run. Together, these findings tell us that actual wage bargaining outcomes for central government workers are in line with the stated intentions of the Swedish Agency for Government Employers and that they are not placing undue pressure on the private sector market for white-collar workers.

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A Pre-Testing the Data

In this Appendix, the variables w_t^{ps} and w_t^{cg} are subjected to a series of tests to determine their order of integration and to test for the presence of deterministic trends. This is done using the augmented Dickey-Fuller sequential procedure as outline in Enders (2004).

A.1 Testing w_t^{ps} for Stationarity

1. Run a Dickey-Fuller regression for the variable w_t^{ps} with a constant and a time trend

$$\Delta w_t^{ps} = \underset{(78.63)}{300} + \underset{(28.27)}{63.1} t - \underset{(0.038)}{0.05} w_{t-1}^{ps} + e_t.$$

The Q-statistics tell us that there is no serial correlation in e_t . Hence, the DF-test is a valid regression.

- (a) H_0 : Coefficient on $w_{t-1}^{ps} = 0 \rightarrow$ unit root.
 - (b) H_A : Coefficient on $w_{t-1}^{ps} < 0 \rightarrow$ no unit root.
 - i. Test statistic = $-0.05/0.038 = -1.34$.
 - ii. MacKinnon critical value = -3.21 (10% level).
 - (c) Test result = we cannot reject the presence of a unit root.
2. Test for the presence of the trend.
 - (a) H_0 : Coefficients on $w_{t-1}^{ps} = t = 0$.
 - (b) H_A : Coefficients on $w_{t-1}^{ps} = 0$ and $t \neq 0$.
 - i. F -test using ϕ_3 test statistic (Enders, 2004, p. 440)
 - ii. F -statistic = 20.06
 - iii. Critical value = 10.61 (1% significance level).
 - (c) Test result = reject H_0 .
 3. Test for unit root using normal distribution.

- (a) H_0 : Coefficient on $w_{t-1}^{ps} = 0$.
 - (b) H_1 : Coefficient on $w_{t-1}^{ps} < 0$.
 - i. t -value = -1.34
 - ii. We cannot reject H_0 .
4. Test results = w_t^{ps} contains both a stochastic trend and a deterministic trend.
5. Test for a second unit root.
6. Run a Dickey-Fuller regression for the variable Δw_t^{ps} with a constant and a time trend

$$\Delta \Delta w_t^{ps} = \frac{282}{(100.6)} - \frac{21.3}{(6.616)} t - \frac{0.86}{(0.186)} \Delta w_{t-1}^{ps} + e_t.$$

The Q-statistics tells us that there is no serial correlation in e_t . Hence, the DF-test is valid.

- (a) H_0 : Coefficient on $\Delta w_{t-1}^{ps} = 0 \rightarrow$ second unit root.
 - (b) H_1 : Coefficient on $\Delta w_{t-1}^{ps} < 0 \rightarrow$ no second unit root.
 - i. Test statistic = $-0.86/0.186 = -4.61$.
 - ii. MacKinnon critical value = -4.28 (1% level).
 - (c) Test result = we can reject the presence of a second unit root.
7. Is the coefficient on t significant? Yes \rightarrow quadratic trend.
8. We conclude that w_t^{ps} is an $I(1)$ variable. It also contains a quadratic, deterministic trend.

A.2 Testing w_t^{cg} for Stationarity

1. Run a Dickey-Fuller regression for the variable w_t^{cg} with a constant and a time trend

$$\Delta w_t^{cg} = \frac{300}{(80.08)} + \frac{79.2}{(29.75)} t - \frac{0.084}{(0.043)} w_{t-1}^{cg} + e_t.$$

The Q-statistics tell us that there is no serial correlation in e_t . Hence, the DF-test is valid.

- (a) H_0 : Coefficient on $w_{t-1}^{cg} = 0 \rightarrow$ unit root.
- (b) H_A : Coefficient on $w_{t-1}^{cg} < 0 \rightarrow$ no unit root.
 - i. Test statistic = $-0.084/0.043 = -1.94$.
 - ii. MacKinnon critical value = -3.21 (10% level).
- (c) Test result = we cannot reject the presence of a unit root.

2. Test for the presence of the trend.
 - (a) H_0 : Coefficients on $w_{t-1}^{cg} = t = 0$.
 - (b) H_A : Coefficients on $w_{t-1}^{cg} = 0$ and $t \neq 0$.
 - i. F -test using ϕ_3 test statistic (Enders, 2004, p. 440).
 - ii. F -statistic = 15.36
 - iii. Critical value = 10.61 (1% significance level).
 - iv. Test result = reject H_0 .
3. Test for unit root using normal distribution.
 - (a) H_0 : Coefficient on $w_{t-1}^{cg} = 0$.
 - (b) H_A : Coefficient on $w_{t-1}^{cg} < 0$.
 - i. t -value = -1.94.
 - ii. We cannot reject H_0 at 5% significance levels (p -value = 0.062).
4. Test results = w_t^{cg} contains both a stochastic trend and a deterministic trend.
5. Test for a second unit root.
6. Run a Dickey-Fuller regression for the variable Δw_t^{cg} with a constant and a time trend

$$\Delta \Delta w_t^{cg} = \frac{255}{(103.0)} - \frac{16.4}{(6.246)} t - \frac{0.78}{(0.184)} \Delta w_{t-1}^{cg} + e_t.$$

The Q-statistics tells us that there is no serial correlation in e_t . Hence, the DF-test is valid.

 - (a) H_0 : Coefficient on $\Delta w_{t-1}^{cg} = 0 \rightarrow$ unit root.
 - (b) H_A : Coefficient on $\Delta w_{t-1}^{cg} < 0 \rightarrow$ no unit root.
 - i. Test statistic = $-0.78/0.184 = -4.23$.
 - ii. MacKinnon critical value = -4.28 (1% level)
 - (c) Test result = we can reject the presence of a unit root in $\Delta \Delta w_t^{cg}$.
7. Is the coefficient on t significant? Yes \rightarrow quadratic trend.
8. We conclude that w_t^{cg} is an $I(1)$ variable. It also contains a quadratic, deterministic trend.