Applied Economics, 2006, **38**, 1617–1625

Is the Swedish central government a wage leader?

Matthew J. Lindquist[a,*] and Roger Vilhelmsson[b]

[a]*Swedish Institute for Social Research, Stockholm University, S-106 91 Stockholm, Sweden*
[b]*Swedish Agency for Government Employers, 103 65 Stockholm, Sweden*

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. It is first shown that salaries of white-collar workers in the private sector and central government are cointegrated. It is then found that private sector salaries are weakly exogenous to the system of equations. This means that the private sector is the wage leader in the long-run model. It is also found 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.

I. Introduction

The Scandinavian countries have small, open economies. Their labour 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-à-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 1970s. The Norwegian multi-sector price income model (Aukrust, 1970, 1977), the Swedish EFO-model (Edgren *et al*., 1973), and Finland's input-output framework (Halttunen and Molander, 1972) all address this problem explicitly.[2] 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 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).[3]

*Corresponding author. E-mail: matthew.lindquist@sofi.su.se

[1] If 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.

[2] See also Jacobsson and Lindbeck (1971) for related work on inter-sectoral wage linkages.

[3] Wage leadership can also be derived from wage bargaining models, efficiency wage models (see e.g. Bemmels and Zaidi, 1990), models of impeded migration (Latreille and Manning, 2000) and market models with productivity leaders (see Section II). The phrase wage leadership has come to dominate the Scandinavian debate. Its intended use is similar to the term wage spillovers often found in the international literature (see e.g. McGuire and Rapping, 1968; Addison and Burton, 1979; Breitung and Meyer, 1994; Budd, 1997; Latreille and Manning, 2000). The definition of wage leadership is also related to Budd's (1997) concept of pattern bargaining and McGuire and Rapping's (1968) notion of a key bargain.

Applied Economics ISSN 0003–6846 print/ISSN 1466–4283 online
http://www.tandf.co.uk/journals
DOI: 10.1080/00036840500407124

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 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 study is to examine whether or not actual wage outcomes of central government employees are in line with this stated praxis.[4]

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 Fig. 1). Fourth, in 2002, the central government employed only 6% of all workers, while local government employed 28.5% and the private sector employed the remaining 65.5%.[5] Together, these facts makes it less likely that the central government has been acting as a wage leader.[6]

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. However, two 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). The present study challenges these results and re-establishes the fact that there is no wage push coming from the central government.[7]

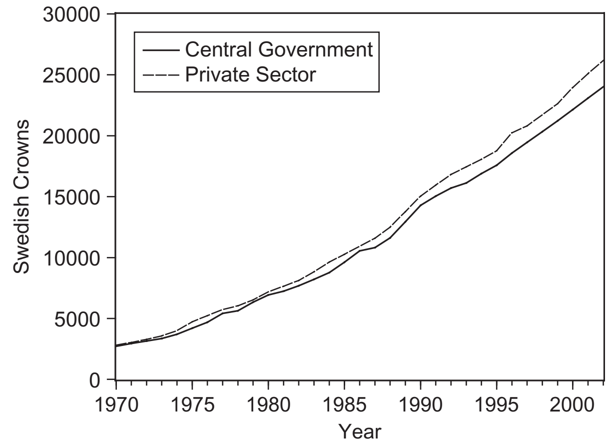

Fig. 1. Monthly Salaries of White-Collar Workers, 1970–2002

To this end, a simple, two-sector model of wage determination and wage leadership is first introduced in Section II. An empirical version of this model is presented in Section III along with a formal definition of wage leadership and a brief description of our empirical tests.

These tests are carried out using a unique, high quality data set, which is presented in Section IV. Unlike the previous studies by Holmlund and Ohlsson (1992), Jacobson and Ohlsson (1994), Tägtström (2000) and Friberg (2003), the salaries of central government employees are not compared with the salaries of all workers in the private sector (i.e., an aggregate of blue-collar and white collar workers). Instead, they are compared with white collar salaries alone. This is believed to be the proper comparison, since more than 95% of all central government employees are white-collar workers covered by white-collar contracts (Lindquist and Vilhelmsson, 2004). If the central government is putting undue pressure on the labour market, this effect should

[4] It is not the goal of this study to analyse the validity of the normative conclusions drawn from the EFO-model. These are taken as praxis. The goal is to put this praxis to a rigorous statistical test.

[5] Central government employment peaked at 11.5% 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).

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

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

be most noticeable in the market for white-collar workers.[8]

The empirical results of this study are presented in Section V. 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 study are similar to those used by Jacobson and Ohlsson (1994) and Breitung and Meyer (1994).

There are two primary results. First, private sector salaries are found to be weakly exogenous to the cointegrated 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 is interpreted in line with the EFO-model as the sum of changes in exogenous productivity and changes in exogenous world prices). Together, these findings tell one 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.

II. A Two-Sector Model of Wage Determination and Wage Leadership

This section develops a simple, two-sector model of relative wage determination. To this end, it is assumed that there is a representative firm in each sector that maximizes profits

$$\Pi_{i,t} = P_{i,t} y_{i,t} - W_{i,t} l_{i,t} \tag{1}$$

subject to a Cobb-Douglas production function

$$y_{i,t} = a_{i,t} (l_{i,t})^{\theta_i} \quad 0 < \theta_i < 1 \tag{2}$$

and to the law of motion for sector-specific technological change

$$a_{i,t} = \alpha_i + \rho a_{i,t-1} + \gamma_i t + \varepsilon_{i,t}$$
$$0 < \rho < 1, \ \gamma_i > 0, \ \varepsilon_{i,t} \sim N(0, \sigma_i) \tag{3}$$

where $P_{i,t}$ is the price of good $y_{i,t}$ produced in sector $i \in \{1, 2\}$ at time t and $W_{i,t}$ is the nominal remuneration to labour, $l_{i,t}$, and where sector-specific technology, $a_{i,t}$, follows a trend stationary $AR(1)$ process. Equilibrium wages, $W^*_{i,t}$, can be derived from the first order conditions of the profit maximizing firm under the assumption of perfectly competitive goods and factor markets

$$\frac{\partial \Pi_{i,t}}{\partial l_{i,t}} : P_{i,t} \theta_i a_{i,t} (l_{i,t})^{\theta_i - 1} - W^*_{i,t} = 0. \tag{4}$$

Imposing a market clearing condition gives us

$$W^*_{1,t} = W^*_{2,t}. \tag{5}$$

If labour is not perfectly mobile between the two sectors, then sector-specific productivity shocks, ε_{it}, will temporarily disrupt the equilibrium wage equation. But over time (i.e., on average), Equation 5 must hold. Thus, writing Equation 5 in vector error-correction (VEC) form is a natural reformulation of the equilibrium relationship between wages in our two sector model

$$\Delta \mathbf{W}_t = \boldsymbol{\pi} \mathbf{W}_{t-1} + \boldsymbol{\pi}_1 \Delta \mathbf{W}_{t-1} + \cdots + \boldsymbol{\pi}_p \Delta \mathbf{W}_{t-p} + \varepsilon_t \tag{6}$$

where $\mathbf{W}_t = \left[W_{1,t} \ W_{2,t} \right]'$, each $\boldsymbol{\pi}$ is a (2×2) matrix of model coefficients and ε_t is a (2×1) vector of sector-specific productivity shocks.[9]

By rewriting Equation 5 in VEC form we have not defined *ex ante* which of the two sectors (if any) is the wage (productivity) leader, where the concept of wage leadership is defined in terms of Granger causality.[10] Furthermore, by allowing for up to p lags, we have not specified the speed of the adjustment process. Both of these questions are empirical in nature and will be determined in Section V by submitting the empirical model to a series of econometric tests.

McGuire and Rapping (1968) extended their simple, market model of wage determination to include a 'key' bargain in the wage leading sector,

[8] Andersson and Isaksson (1997) also make this distinction between white-collar and blue-collar workers, but these data set only goes up to 1995. The present updated data set allows one 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.

[9] The first difference of x is denoted by Δx.

[10] This is the most commonly used definition of wage leadership found in the literature, see, e.g., Bemmels and Zaidi (1990), Holmlund and Ohlsson (1992), Jacobson and Ohlsson (1994), Mizala and Romaguera (1995), Andersson and Isaksson (1997), Tägtström (2000), Friberg (2003).

the results of which 'spill over' into other sectors, albeit with some delay

$$W_{2,t} = W_{2,t-1}\left[\frac{\left(W^{*}_{2,t}\right)^{\delta_2}\left(W_{1,t}\right)^{1-\delta_2}}{W_{2,t-1}}\right]^{\delta_1}$$

$$0 < \delta_1 < 1, \quad 0 < \delta_2 < 1 \tag{7}$$

The market wage, $W^{*}_{2,t}$, is the same as before. But now, the wage in sector 2, $W_{2,t}$, is determined by past wages in that sector as well as an adjustment term which is influenced by both market forces (represented by the market wage, $W^{*}_{2,t}$) and by the key bargain struck in the wage leading sector (represented by the wage in sector 1, $W_{1,t}$).

It is this type of bargaining mechanism that most authors have in mind when discussing wage leadership in Sweden. The Swedish EFO-model states that it is the competitive sector that strikes the key bargain and acts as the wage leader. The study has chosen, however, to remain agnostic as to which sector is the wage leader when formulating the empirical model.

III. An Empirical Model of Wage Leadership

Equation 6 can be viewed as the reduced form version of Equation 7. It is Equation 6 that is taken into the data. To do this, the following statistical version of Equation 6 is posited:

$$\Delta\mathbf{W}_t = \boldsymbol{\mu} + \boldsymbol{\pi}\mathbf{W}_{t-1} + \boldsymbol{\pi}_1\Delta\mathbf{W}_{t-1} + \cdots + \boldsymbol{\pi}_p\Delta\mathbf{W}_{t-p} + \varepsilon_t \tag{8}$$

where $\mathbf{W}_t = \left[W_{1,t}\ W_{2,t}\right]'$, $\boldsymbol{\mu}$ is a (2×1) vector of potentially nonzero constants, each $\boldsymbol{\pi}$ is a (2×2) matrix of regression coefficients and ε_t is a (2×1) vector of Gaussian white noise error terms.

Addison and Burton (1979) pointed out the difficulties in identifying the cause and/or source of wage spillovers. Are they due to market forces, institutions or equity considerations? The VEC(p) model allows one to circumvent this identification problem and to remain agnostic about the underlying causal mechanisms. This study simply addresses the questions of existence and direction of wage spillover effects.

The VEC modelling strategy allows one to identify the wage leader and to test for wage leadership in two distinct ways. First, one can examine if one of the wage variables included in the model is weakly exogenous to the estimated system of equations. If nominal wages in sectors 1 and 2 are cointegrated (as predicted by the present), and if the nominal wage in sector 1 turns out to be weakly exogenous, while the nominal wage in sector 2 is not, then one knows that $W_{2,t}$ (alone) adjusts to changes $W_{1,t}$ in order to maintain the long-run equilibrium. In this case, sector 1 is deemed the wage leading sector and $W_{1,t}$ is the key bargain or reference wage. Sector 2 is the wage follower. Second, the model allows one to construct a robust test of Granger causality between $\Delta W_{1,t}$ and $\Delta W_{2,t}$. One which does not suffer from the exclusion of a very important variable, namely, the long-run cointegrating relationship between $W_{1,t}$ and $W_{2,t}$.

The following definitions are now proposed:

Definition 1 (Wage Leadership in the Long-Run Model) *Sector 1 is the wage leader in the long-run model represented by* $\boldsymbol{\pi}_0\mathbf{W}_{t-1}$ *in Equation* 8 *if the variables* $\mathbf{W}_{t-1}$ *are stationary, trend stationary or cointegrated and if* $\mathbf{W}_{1,t-1}$ *is found to be weakly exogenous to the system of equations and* $\mathbf{W}_{2,t-1}$ *is shown to be endogenous to the system of equations.*

Definition 2 (Wage Leadership in the Short-Run Model) *Sector* 1 *is the wage leader in the short-run model represented by* $\Delta\mathbf{W} = \boldsymbol{\pi}_1\Delta\mathbf{W}_{t-1} + \cdots + \boldsymbol{\pi}_p\mathbf{W}_{t-p} + \varepsilon_t$ *in Equation* 8 *if lagged values of* $\Delta\mathbf{W}_{1,t}$ *can be used to predict current values of* $\Delta\mathbf{W}_{2,t}$ *and lagged values of* $\Delta\mathbf{W}_{2,t}$ *cannot be used to predict current values of* $\Delta\mathbf{W}_{1,t}$.

IV. Data

Two data series are used in this study: nominal monthly, white-collar salaries in the private sector, $W_{ps,t}$, and nominal monthly salaries in the central government, $W_{cg,t}$ (see Fig. 1). The data are annual time series from 1970 to 2002 collected by the Confederation of Swedish Enterprise (Svenskt Näringsliv) and the Swedish Agency for Government Employers (Arbetsgivarverket). They are based on actual contracts and cover nearly all workers in these two categories.

There are three major advantages of using these data, as compared with the data used in earlier studies. First, and most importantly, since more than 95% of all central government workers are white-collar workers, and since all workers are covered by white-collar 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.[11]

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 they are treated as today's wages. This type of periodicization is made possible by the fact that the data come 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.

Examining the time series for private sector, white-collar salaries and central government salaries in Fig. 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, each variable is pre-tested 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).[12]

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 1970s and 1980s. 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 one that one 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.

V. Estimating the Vector Error-Correction Model

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 in levels.[13] The VAR(p) model can be written as

$$\mathbf{W}_t = \boldsymbol{\mu} + \mathbf{A}_1\mathbf{W}_{t-1} + \mathbf{A}_2\mathbf{W}_{t-2} + \cdots + \mathbf{A}_p\mathbf{W}_{t-p} + \varepsilon_t \tag{9}$$

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

Multivariate generalizations of the Akaike information criterion (AIC) and Bayesian information criterion (BIC) are used to choose the appropriate lag length. The principle behind these two tests is the same. One is 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 one that a VAR($p > 1$) model does not significantly outperform the VAR(1) model. The results of these tests are reported in Table 1.

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

[11] See Latreille and Manning (2000) for more on inter-occupational wage spillovers.
[12] For more detailed information concerning these tests and their results see Lindquist and Vilhelmsson (2004).
[13] Since 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.

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

Table 2. Determining the lag length, *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

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 have not yet been extracted from the residuals.

The study continues by estimating a VAR(2) model. The residuals from this model are normally distributed, they do not suffer from serial autocorrelation nor is the presence of ARCH detected. 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

$$\mathbf{W}_t = \boldsymbol{\mu} + \mathbf{A}_1\mathbf{W}_{t-1} + \mathbf{A}_2\mathbf{W}_{t-2} + \varepsilon_t \tag{10}$$

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

$$\Delta\mathbf{W}_t = \boldsymbol{\mu} + \boldsymbol{\pi}\mathbf{W}_{t-1} + \boldsymbol{\pi}_1\Delta\mathbf{W}_{t-1} + \varepsilon_t. \tag{11}$$

Testing for cointegration between the nonstationary variables, $\mathbf{W}_{t-1}$, 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, one wants 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

$$\begin{aligned}\Delta\mathbf{W}_t &= \boldsymbol{\mu}_{sr} + \boldsymbol{\delta}_{sr}t + \tilde{\boldsymbol{\pi}}\tilde{\mathbf{W}}_{t-1} + \boldsymbol{\pi}_1\Delta\mathbf{W}_{t-1} + \varepsilon_t \\ &= \boldsymbol{\mu}_{sr} + \boldsymbol{\delta}_{sr}t + \underbrace{\boldsymbol{\alpha}\begin{bmatrix}\beta_{ps}\\ \beta_{cg}\\ \mu_{lr}\\ \delta_{lr}\end{bmatrix}'}_{\tilde{\boldsymbol{\pi}}}\underbrace{\begin{bmatrix}W_{ps,t-1}\\ W_{cg,t-1}\\ 1\\ t\end{bmatrix}}_{\tilde{\mathbf{W}}_t} \\ &\quad + \boldsymbol{\pi}_1\Delta\mathbf{W}_{t-1} + \varepsilon_t\end{aligned} \tag{12}$$

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 $\mathbf{W}_{t-1}$ 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 $\tilde{\mathbf{W}}_{t-1}$, 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 five 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.

One's task is to identify which of these models fits the data best at the same time as they are tested for cointegration. This can be done by testing different sets of restrictions jointly with the restriction that the rank of $\tilde{\boldsymbol{\pi}}$ is either 0, 1, or 2. one can minimize on the number of tests necessary to complete this task by realizing that neither Model 1 nor Model 2 are reasonable representations 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

Table 3. Joint determination of rank($\widetilde{\pi}$) and deterministic components

H_0 : rank($\widetilde{\pi}$)	Model 3 LR-stat	Model 3 5%	Model 4 LR-stat	Model 4 5%	Model 5 LR-stat	Model 5 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

Note: * denotes rejection of H_0 at 5% significance level. Critical values taken from Osterwald-Lenum (1992).

what is found when the variables were pre-tested. One can also exclude the test for rank($\widetilde{\boldsymbol{\pi}}$) = 2, since both variables are $I(1)$.

This leaves one with a set of six 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 $3 \cap \text{rank}(\widetilde{\pi}) = 0$; model $4 \cap \text{rank}(\widetilde{\pi}) = 0$; model $5 \cap \text{rank}(\widetilde{\pi}) = 0$; model $3 \cap \text{rank}(\widetilde{\pi}) = 1$; model $4 \cap \text{rank}(\widetilde{\pi}) = 1$; model $5 \cap \text{rank}(\widetilde{\pi}) = 1$. Table 3 shows each of these null hypotheses along with the appropriate likelihood-ratio (trace) test.

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

H_0: model $5 \cap \text{rank}(\widetilde{\pi}) = 1$ 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 the preferred model.

The rank of the estimated $\widetilde{\boldsymbol{\pi}}$ matrix, $\widehat{\boldsymbol{\pi}}$, is equal to one which means that the long-run model $\widetilde{\boldsymbol{\pi}}\widetilde{\mathbf{W}}_{t-1}$ 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 1970s and 1980s.

Testing for wage leadership

The hypotheses concerning wage leadership can be formulated as restrictions on the VEC(1) model and then tested. To do this, the (2×4) matrix $\widehat{\boldsymbol{\pi}}$ is factored into a (2×1) vector, $\widehat{\boldsymbol{\alpha}}$, and a (4×1) vector, $\widehat{\boldsymbol{\beta}}$, such that $\widehat{\boldsymbol{\pi}} = \widehat{\boldsymbol{\alpha}}\widehat{\boldsymbol{\beta}}'$. The first vector, $\widehat{\boldsymbol{\alpha}} = \left[\widehat{\alpha}_{ps}\ \widehat{\alpha}_{cg}\right]$, 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 $\mathbf{W}_t = \left[W_{ps,t}\ W_{cg,t}\right]'$. The coefficients in $\widehat{\boldsymbol{\alpha}}$ measure the speed of adjustment to past equilibrium errors. The second vector, $\widehat{\boldsymbol{\beta}}' = [\widehat{\beta}_{ps}\ \widehat{\beta}_{cg}\ \widehat{\mu}_{lr}\ \widehat{\delta}_{lr}]$, is the cointegrating vector itself, which defines the long run equilibrium relationship between $W_{ps,t-1}$ and $W_{cg,t-1}$.

Table 4. Estimates of $\widehat{\boldsymbol{\alpha}}$ and $\widehat{\boldsymbol{\beta}}$

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

Note: [a] Standard errors in parentheses.

The estimated values of $\widehat{\boldsymbol{\alpha}}$ and $\widehat{\boldsymbol{\beta}}$ are reported in Table 4. The values of $\widehat{\boldsymbol{\beta}}$ have been normalized with $\widehat{\beta}_{ps}$. The standard errors of those coefficients in $\widehat{\boldsymbol{\beta}}$ that are not uniquely identified are not reported.

Testing for wage leadership in the long-run model amounts to testing each variable for weak exogeneity (recall Definition 1). The existence of only one cointegrating vector simplifies this test: one need only examine the t-values associated with $\widehat{\alpha}_{ps}$ and $\widehat{\alpha}_{cg}$. These are −0.044 and 2.96, respectively. Since we cannot reject $H_0 : \alpha_{ps} = 0$, we conclude that $W_{ps,t-1}$ is weakly exogenous to the system of equations. Central government salaries, $W_{cg,t-1}$, on the other hand, are endogenous to the system of equations, since one can clearly reject $H_0 : \alpha_{cg} = 0$.

The null hypothesis that the central government acts as a wage leader $(H_0 : \widehat{\alpha}_{ps} \neq 0 \cap \widehat{\alpha}_{cg} = 0)$ is strongly rejected. On the other hand, one cannot reject the hypothesis that the private sector is the wage leader $\left(H_0 : \widehat{\alpha}_{ps} = 0 \cap \widehat{\alpha}_{cg} \neq 0\right)$. The test for weak exogeneity shows 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, $\widehat{\alpha}_{cg} = 0.769$ tells us that the central government corrects 77% 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.

Given that $W_{ps,t-1}$ is weakly exogenous to the system of equations, one can factorize the model into two single equations: one marginal model of $\Delta W_{ps,t}$ and one conditional model for $\Delta W_{cg,t}$.[14]

[14] This also means that one can estimate each equation separately using OLS.

One can use the marginal model of $\Delta W_{ps,t}$ to test whether or not $\Delta W_{cg,t}$ Granger causes $\Delta W_{ps,t}$. This can be viewed as a test for wage leadership in the short-run model (recall Definition 2).

Estimating the marginal model of $\Delta W_{ps,t}$ results in the following regression equation

$$\Delta W_{ps,t} = \underset{(103.5)}{267} + \underset{(6.771)}{20.5t} + \underset{(0.245)}{0.02}\,\Delta W_{ps,t} + \underset{(0.233)}{0.17}\,\Delta W_{cg,t} + \varepsilon_t^{ps}. \quad (13)$$

An F-test concerning the hypothesis that the two coefficients in $\boldsymbol{\pi}_1$ are equal to zero has a p-value of 0.58. So, one can pare down the marginal model to

$$\Delta W_{ps,t} = \underset{(103.5)}{267} + \underset{(6.771)}{20.5t} + \varepsilon_t^{ps} \quad (14)$$

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

VI. Conclusion

This study 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 Bank (Tägtsröm, 2000 and Friberg, 2003). The findings tell one 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.

Acknowledgements

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

References

Addison, J. and Burton, J. (1979) The identification of market and spillover forces in wage inflation: a cautionary note, *Applied Economics*, **11**, 95–104.

Andersson, P. and Isaksson, A. (1997) Is private sector pay formation affected by the central government in Sweden: a causal analysis, mimeograph, Swedish Agency for Government Employers, Stockholm, Sweden.

Aukrust, O. (1970) PRIM I: a model of the price income distribution mechanism of an open economy, *Review of Income and wealth*, **16**, 51–78.

Aukrust, O. (1977) Inflation in the Open Economy: a Norwegian model, Article series Nr. 96, Statistisk Sentralbyrå, Oslo, Norway.

Bemmels, B. G. and Zaidi, M. A. (1990) Wage leadership in Canadian industry, *Applied Economics*, **22**, 553–67.

Breitung, J. and Meyer, W. (1994) Testing for unit roots in panel data: are wages on different bargaining levels cointegrated?, *Applied Economics*, **26**, 353–61.

Budd, J. W. (1997) Institutional and market determinants of wage spillovers: evidence from UAW pattern bargaining, *Industrial Relations*, **36**, 97–116.

Edgren, G., Faxén, K.-O. and Odhner, C.-E. (1973) *Wage Formation and the Economy*, Allen & Unwin, London, England.

Elvander, N. (2004) Från plan till marknad. Statlig lönebildning 1989–2003 (From Plan to Market. Central Government Wage Formation 1989–2003), *Ekonomisk Debatt 2004*, **1**, 5–21.

Enders, W. (2004) *Applied Econometric Time Series*, 2nd Edn, John Wiley & Sons, Inc., Hoboken, NJ, USA.

Friberg, K. (2003) Intersectoral wage linkages in Sweden, Sveriges Riksbank Working Paper Series Nr. 158, Stockholm, Sweden.

Halttunen, H. and Molander, A. (1972) The input-output framework as a part of a macroeconomic model: production-price-income block in the Bank of Finland quarterly econometric model, *Finnish Economic Journal*, **68**, 219–39.

Holmlund, B. and Ohlsson, H. (1992) Wage linkages between private and public sectors in Sweden, *Labour*, **6**, 3–17.

Jacobson, T. and Ohlsson, H. (1994) Long-run relations between private and public sector wages in Sweden, *Empirical Economics*, **19**, 343–60.

Jacobsson, L. and Lindbeck, A. (1971) On the transmission mechanism of wage change, *Swedish Journal of Economics*, **73**, 273–93.

Johansen, S. (1995) *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Advanced Texts in Econometrics, Oxford University Press, Oxford, England.

Latreille, P. L. and Manning, N. (2000) Inter-industry and inter-occupational wage spillovers in UK manufacturing, *Oxford Bulletin of Economics and Statistics*, **62**, 83–99.

Lindquist, M. J. and Vilhelmsson, R. (2004) Är staten löneledande? En ekonometrisk studie av de statliga anstä llda och de privata tjänstemännen 1970–2002 (Is the Central Government a Wage Leader? An Econometric Study of Central Government Employees and Private Sector Employees 1970–2002), mimeograph, Swedish Agency for Government Employers, Stockholm, Sweden.

McGuire, T. W. and Rapping, L. A. (1968) The role of market variables and key bargains in the manufacturing wage determination process, *Journal of Political Economy*, **76**, 1015–36.

Mizala, A. and Romaguera, P. (1995) Testing for wage leadership processes in the Chilean economy, *Applied Economics*, **27**, 303–10.

Osterwald-Lenum, M. (1992) A note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics, *Oxford Bulletin of Economics and Statistics*, **54**, 461–72.

Sims, C. (1980) Macroeconomics and reality, *Econometrica*, **48**, 1–49.

Tägtström, S. (2000) The wage spread between different sectors in Sweden, *Sveriges Riksbank Economic Review 2000*, **4**, 77–82.