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# Wage Formation and the Relation between Real Wages and Unemployment in Sweden

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## Abstract

This paper examines the determinants of the wage level and the relation between the wage level and unemployment in Sweden between 1982 and 2002, using a cointegrated VAR approach. The long-run relation between wages and unemployment is found to be negative. There is also evidence of large deviations between the development in real wages and the development in productivity, in that the growth in real wages exceeds the growth in labour productivity. The results indicate that rigidities are present in the Swedish labour market and that these rigidities may cause higher unemployment. Furthermore, trade unions are found to have a large influence on the wage formation process.

**JEL Classification:** E24, C32

**Keywords:** Wage formation, Unemployment, Cointegration, Common Trends

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# 1 Introduction

This paper is concerned with wage formation and the labour market in Sweden between 1982 and 2002, a period in which the problem of unemployment became severe in Sweden. A large part of current domestic policy is devoted to handling the unemployment issue. To be able to take appropriate action against high unemployment, it is vital to have firm knowledge about the origins of the unemployment and the factors that influence its persistence. Therefore, identifying the sources of imbalances in the labour market and analysing the mechanisms that restore equilibrium are of great importance for economic policy decisions.

In the 1980s, Sweden suffered from large price inflation and large wage inflation, problems that are often argued to stem from the prevailing wage determination process. Economic theory predicts nominal wages to be equal to the value of the marginal product of labour, implying that nominal wage increases correspond to the sum of the increases in labour productivity and price inflation. Therefore, we emphasize the interrelation of wages and productivity and prices in the empirical analysis.

We also analyse the role of trade unions in the wage formation process. Both employers and employees are concerned with the real wage rather than the nominal wage, but employers are concerned with the real wage in terms of producer prices, while employees are concerned with the real wage in terms of consumer prices. An interesting question is, therefore, which of these prices are most important in determining the real wage. If consumer prices are more important in determining the wage level, we interpret this as a high degree of trade union influence in the wage formation process, but if producer prices are more important, firms have the stronger position.

When designing an appropriate policy concerning unemployment, it is essential to possess information about the different types of adjustment processes that are present on the labour market. This information may enable us to design a policy that speeds up those adjustment processes, or to act against inflexibilities in the labour market. Here, we study the adjustment process in both wages and unemployment as a function of different types of deviations from equilibrium.

We use cointegration analysis to estimate the wage formation in a model including wage and price setting behaviour. In an examination of the relations between wages and unemployment, it is important to include the interaction between wages and prices, since expectations of future prices play a vital role in the wage formation process. In accordance with the above discussion regarding the role of trade unions, we include both domestic producer and consumer prices in the analysis. Since Sweden is a small, open economy with a large dependency on trade, we also include import prices to account for the foreign dependency. The final model thus includes labour market variables, price variables and a foreign variable. Using cointegration analysis, we estimate the long-run equilibrium relations connecting wages, prices and unemployment to each other and the short-run adjustment towards these steady state relations. We also examine the factors that influence the trends in wages and unemployment by performing a common trends analysis.

We find that the elasticity between the real wage and unemployment is negative, so when unemployment increases, there is a downward pressure on the real wage level. Furthermore, the estimates indicate that the growth in real wages exceeds the growth in labour productivity throughout the studied period, implying that the wage increases are too large compared to productivity growth. Moreover, the development in real wages is to a large extent influenced by the development of consumer prices, a feature interpreted to be the result of a high degree of trade union influence in the wage formation process. In the common trends analysis, we find that unemployment does not react strongly to domestic productivity shocks. Instead, unemployment reacts strongly to shocks in the labour market and to changes in the monopoly situation in the producer market. We interpret this as signs of rigidities in the Swedish labour market since, without rigidities, these shocks should have resulted in changes in wages instead of in unemployment.

The paper is organized as follows. Previous studies are discussed in section 2. Section 3 discusses the theoretical framework for the empirical analysis. Preliminary results regarding the empirical cointegration model are presented in section 4. Section 5 contains the result of the cointegration analysis and the analysis of the estimated wage setting behaviour and price behaviour. The underlying common trends are identified and interpreted

economically in a structural common trends model, and the results of the common trends analysis and an analysis of impulse response functions are included in section 6. Section 7 concludes this study.

## 2 Previous studies

The relation between wages and unemployment has been analysed in a number of studies, both theoretically and empirically. The general conclusion is that the relation between the variables is negative. An overview of these studies can be found in Blanchard and Katz (1999). In OECD (1997), the empirical relation between the growth rate of wages and the unemployment level is estimated for the OECD countries. Besides the negative relationship between wage inflation and unemployment, OECD finds that the wage development in the European countries also depends on previous imbalances between real wages and productivity. This error-correction behaviour is not present in the United States (OECD, 1997). Estimations of wage curves for the Nordic countries, as well as for other European countries, are made in Elmeskov (1994). In a comparison, Elmeskov argues that real wages have reacted more strongly to changes in unemployment in the Nordic countries than in the EU countries. In all of these empirical studies, single equations are estimated where wage inflation is regressed on unemployment.<sup>1</sup>

Pétursson and Sløk (2001) examine wage formation and employment in Denmark. They derive a theoretical model, where trade unions and firms bargain over the real wage, and use this model to estimate the wage formation process using Danish labour market data in a cointegrated VAR setting. Their conclusion is that the main reason for the persistent unemployment in Denmark is wage rigidity, and their estimate of the elasticity of unemployment with respect to real wages is low. Juselius (2003b) looks at the determination of wages, prices, productivity and unemployment in the Euro-area in the post-Bretton Woods-period. Aggregated data for the Euro countries is used and the focus is on analysing the change in inflation and unemployment behaviour that has occurred as a result of increasing European integration. Juselius finds that, in the 1970s, strong labour unions were able

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<sup>1</sup>See Blanchard and Katz (1999) for a discussion about the empirical specifications.

to obtain high wage increases that undermined competitiveness in the Euro-area. Competitiveness was restored by laying off workers, thereby increasing labour productivity. Furthermore, Juselius finds a negative relation between real wages and unemployment. The simultaneous determination of wages and prices is also studied by Bårdsen and Fisher (1999) and Bårdsen et al. (2001) for the United Kingdom and Australia, respectively, using a cointegrated VAR approach. The authors find that wages and prices are jointly determined in both countries. In Australia, high unemployment is negative for wage growth, but the relation is the opposite in the United Kingdom.

Cointegrated VAR models are used to study labour markets in Jacobson et al. (1997), Jacobson et al. (1998) and Hansen and Warne (2001). In these studies, a common trends model identifies the sources of the behaviour in the labour market. Jacobson et al. (1998) investigate, among other questions, the relation between real wages and unemployment in Sweden, finding little support for any relation between the two variables, neither in the short run nor in the long run. We analyse a more recent time period and find a relation between real wages and the unemployment rate. We also work on data with a higher frequency and explicitly include the influence on the Swedish economy from abroad by including import prices and the real exchange rate in the analysis.

In a recent study, using a cointegrated VAR model with quarterly data for Sweden between 1980 and 2003, Friberg (2004) finds a negative link between nominal wages and the unemployment rate and a bi-directional causality between nominal wages and consumer prices. Productivity does not, however, influence the wage determination explicitly, as we would expect from economic theory. All the Scandinavian countries are included in Jacobson et al. (1997), who find evidence of hysteresis in unemployment caused by shocks to technology, labour supply and equilibrium unemployment. The Danish unemployment experience is found in Hansen and Warne (2001), where shocks to the labour supply are the most important source for explaining the behaviour of unemployment.

### 3 The economic model

This section contains a simple theoretical model concerning the determination of wages and prices. The aim is to build a framework for the analysis and to rely on these theoretical relations when setting up an appropriate empirical model.

The wage level depends on productivity and prices, but the conditions in the labour market are also taken into account in the wage bargaining process with the wage depending on unemployment as a measure of labour market tightness. The model used is the same as in Juselius (2003b). A motivation for the wage setting equation is found in Pétursson and Sløk (2001), who derive the behaviour of the real wage in a model with utility-maximizing trade unions, profit-maximizing firms and bargaining over the real wage in markets with monopolistic competition.<sup>2</sup>

Equation (1) describes the determinants of the nominal wage. The nominal wage,  $w_t^{nom}$ , depends on producer prices,  $p_t^y$ , consumer prices,  $p_t^c$ , labour productivity,  $y_t$ , and unemployment,  $u_t$ , according to

$$w_t^{nom} = \gamma_0 + (1 - \gamma_1)p_t^y + \gamma_1 p_t^c + \gamma_2 y_t - \gamma_3 u_t. \quad (1)$$

All variables except the unemployment rate are in natural logarithms. Long-run homogeneity is assumed on the relation between nominal wages, consumer prices and producer prices. Thus, the parameters corresponding to these variables all sum to zero and the parameter  $\gamma_1$  lies in the range of  $0 \leq \gamma_1 \leq 1$ . Increasing labour productivity is expected to induce higher nominal wages and hence,  $\gamma_2 \geq 0$ . Finally, high unemployment is expected to render a downward pressure on the wage level, so  $\gamma_3 > 0$ . Based on the assumption of long-run homogeneity between nominal wages and prices, equation (1) can be rewritten in real terms as

$$w_t^{nom} - p_t^y = \gamma_0 + \gamma_1(p_t^c - p_t^y) + \gamma_2 y_t - \gamma_3 u_t. \quad (2)$$

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<sup>2</sup>The difference between the equations is that in Pétursson and Sløk (2001) unemployment benefits are included in the wage equation to capture the effect of the individual's reservation wage on the real wage. In our model, unemployment benefits are not included. In the empirical analysis, the effect of the reservation wage on the wage setting behaviour is captured either in the constant term (if the reservation wage is constant in the period studied) or in the time trend (if the reservation wage is continuously rising).

In equation (2), the real producer wage is positively related to labour productivity, negatively related to unemployment and positively related to the price wedge between consumer and producer prices. In this specification,  $\gamma_1$  has an additional interpretation as a parameter reflecting the bargaining strength of trade unions compared to the bargaining strength of employers. If  $\gamma_1 = 1$ , the development in nominal wages follows the development in consumer prices while nominal wages follow producer prices if  $\gamma_1 = 0$ . In the former case, trade unions have a strong position in the labour market, getting compensation for every rise in consumer prices, while in the second case, producers have the strong position and tie wage increases to their own cost increases. In the empirical analysis, we use the estimate of  $\gamma_1$  to measure the trade union influence in the wage formation process.

With monopolistic competition in the economy, prices are set as mark-ups over production costs. In this theoretical framework, the firms' production costs consist of the nominal wage and the price of imported goods, denoted  $p_t^*$ , and hence, producer prices are set as

$$p_t^y = \theta_0 + \theta_1 w_t^{nom} + (1 - \theta_1) p_t^*. \quad (3)$$

$\theta_1$  and  $(1 - \theta_1)$  measure the weights by which nominal wages and import prices influence the producer price. The mark-up pricing resulting from monopolistic competition is captured in the parameter  $\theta_0$ .<sup>3</sup> Finally, we assume that monopolistic competition also prevails in the final distribution line and specify consumer prices as  $p_t^c = \theta_2 + p_t^y$ .  $\theta_2$  is a measure of the elasticity of substitution between consumer goods that can also be interpreted as a measure of firms' degree of monopoly in the market for consumer goods. The variables in the price equations are in logarithms.

## 4 Long-run homogeneity and the question of I(2)

The theoretical relations in the previous section serve as guidelines in the formulation of the empirical model. The empirical analysis is conducted in a cointegrated VAR model, an

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<sup>3</sup>When deriving equation (3) in a model with monopolistic competition,  $\theta_0$  depends on the weights of nominal wages and import prices in the producer price, and the elasticity of substitution between different goods.

appropriate approach when analysing economic problems with non-stationary data. In the setting of a cointegrated VAR model, we are able to estimate both the long-run relations among the variables, i.e. equations (2) and (3), and the short-run adjustments towards these relations. In the analysis, we use the maximum likelihood approach to cointegration analysis proposed by Johansen (1988, 1991) and Johansen and Juselius (1990, 1992).

Both nominal and real variables are present among the six variables included in the theoretical equations. The inclusion of nominal variables in an empirical analysis can give rise to effects stemming from I(2) behaviour in the data. For example, in empirical studies, price inflation is sometimes found to be I(1), at least in small samples.<sup>4</sup> Inflation being an I(1) variable implies that the price level is I(2), a characteristic that contradicts our usual perceptions of the economy. However, this does not have to be interpreted as an economic characteristic, but instead a statistical property present in an econometric analysis with a limited number of observations. In an econometric analysis, the modelling of the variable improves by treating the variable as having a unit root rather than being stationary.<sup>5</sup> In this section, we conduct an introductory examination of the data and discuss possible signs of I(2) behaviour in the data. Some signs of I(2) are present and we try to identify the sources giving rise to this behaviour by testing for long-run homogeneity between nominal wages and prices.

## 4.1 Data

For the empirical analysis, we use Swedish data between 1982 and 2002 and define a variable vector,  $X_t$ , as

$$X_t = \left[ w_t^{nom} \quad p_t^y \quad p_t^c \quad y_t \quad u_t \quad p_t^* \right]' . \quad (4)$$

Labour productivity,  $y_t$ , is defined as real GDP divided by total employment. All variables except the unemployment rate are in natural logarithms.<sup>6</sup>

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<sup>4</sup>See for example Mishkin (1992), Juselius (1998), Banerjee et al. (2001) and Nielsen (2002).

<sup>5</sup>For a further discussion, see Juselius (1999).

<sup>6</sup> $w^{nom}$  is nominal wages,  $p^y$  is producer prices,  $p^c$  is consumer prices,  $y$  is labour productivity,  $u$  is unemployment and  $p^*$  is import prices. Monthly data on nominal wages, producer prices, consumer prices, total employment and the unemployment rate have been collected from OECD, Main Economic Indicators.

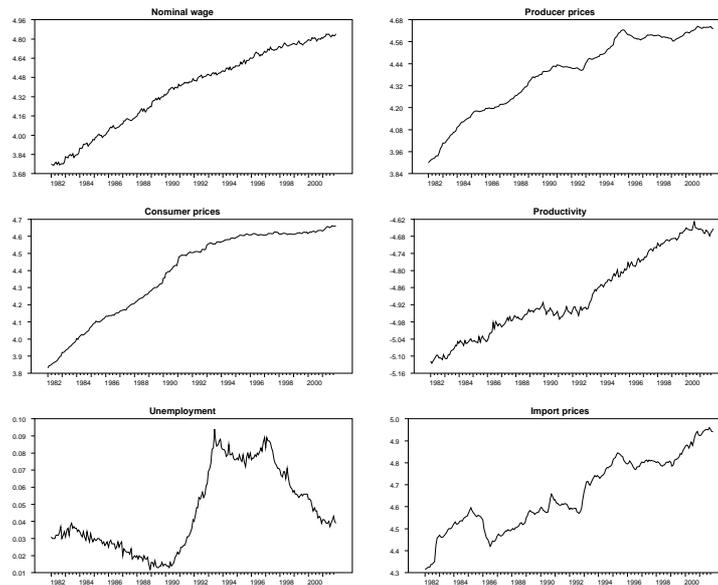


Figure 1: Graphs of nominal wages, producer prices, consumer prices, labour productivity, unemployment and import prices

Figure 1 shows graphs of the six variables. Most of the variables show an upward trend throughout the period. The trend in labour productivity is broken between 1990 and 1993, when Sweden suffered from a deep recession. After 1993, the productivity rises relatively rapidly, a result of large lay-offs in the labour market following the recession and bankruptcies of many firms. The striking feature in the unemployment rate is the large rise in unemployment that occurs in the period. Regarding import prices, upward jumps in the variable can be seen in October 1982 and November 1992, which reflect the devaluation of the Swedish krona in 1982 and the depreciation following the move from a fixed exchange rate to a floating exchange rate policy in 1992.

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The variables are seasonally adjusted by the OECD. Both nominal wages and producer prices are for the manufacturing sector. Monthly data on import prices are from Statistics Sweden, as is nominal GDP. This series is on a quarterly basis and has been linearly interpolated to monthly observations. The series is seasonally adjusted with the X-11 filter in the software EViews. We have tried other interpolation methods, among them a time series interpolation with an ARMA(1,1) model. Overall, this gives very similar result as for the chosen case.

If wages and prices are I(2) and the long-run homogeneity restriction is valid, the vector of nominal variables in equation (4) can be transformed to a vector of real variables that only contains I(1) variables. In this case, the I(2) trend is cancelled. The transformation can be made by deflating nominal wages, producer prices and consumer prices with one of the nominal variables. Under a valid homogeneity restriction, the nominal-to-real transformation can be made without losing any information in the model. Even without a valid restriction, we can apply the transformation, but in this case some information in the model may be lost and a fraction of the I(2) component may still be present in the model. For a thorough discussion about transformations of nominal variables to real variables in cointegration models, see Kongsted (2005). Transforming the variables into real terms simplifies the analysis, since the new data vector only contains I(1) variables and, without the transformation, the analysis should be conducted in an I(2) model.

The first sign of I(2) behaviour in the data can be found by a graphical inspection of the variables in figure 1. We see that nominal wages, consumer prices and, to a lesser extent, producer prices, have a very smooth appearance. Such smoothness may be a sign of the presence of I(2) elements in the data. In order to examine the possibility of I(2) more deeply, we estimate and analyse a VAR model based on the data vector in equation (4).

## 4.2 Empirical specification of the statistical model

The statistical model underlying the empirical analysis is a  $p$ -dimensional vector autoregressive model of order  $k$ , written in vector error correction form as

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \alpha \delta_0 t + \epsilon_t. \quad (5)$$

$X_t$  is a  $(p \times 1)$  variable vector, with  $p$  equal to the number of variables in the model. In our model,  $p = 6$ . The matrices  $\Pi$  and  $\Gamma_i$  are of dimension  $(p \times p)$  and contain autoregressive parameters. It is assumed that  $\epsilon_t$  is independently and identically normally distributed with mean zero and a constant covariance matrix,  $\Omega$ . It is also assumed that the initial values  $(X_{-k+1}, \dots, X_0)$  can be taken as given. The vectors  $\mu$  and  $\alpha \delta_0$  contain deterministic

components in terms of constants and linear drift terms, respectively. In the graphical inspection of the variables in figure 1, several of the variables seemed to contain a linear trend and we therefore allow for linear trends, but no quadratic trends, in the data.  $\mu$  is specified as  $\mu = \alpha\mu_0 + \mu_1$  and left unrestricted in the estimations, implying that  $\mu_0$  measures the intercept in the cointegration relation and  $\mu_1$  the intercept in the first difference in the variables, i.e. the linear growth rate in the data. The time trend,  $t$ , is restricted to the cointegration relation.

Under the presence of cointegration,  $\Pi$  has reduced rank and can be decomposed as  $\Pi = \alpha\beta'$ . The matrices  $\alpha$  and  $\beta$  have the dimension  $(p \times r)$ , with  $r < p$ , where  $r$  is equal to the cointegration rank.  $\beta$  is a matrix of long-run cointegration vectors and the elements in  $\alpha$  are adjustment coefficients, determining the speed of adjustment back to equilibrium after a deviation from the long-run relation.

The appropriate lag length,  $k$ , of equation (5) is selected by the minimization of two information criteria, the Hannan-Quinn criterion and the Schwartz criterion. In panel A in table 1, the values of the two criteria are shown for the inclusion of up to 4 lags. However, the two information criteria give different results. The Hannan-Quinn criterion is minimized for 2 lags whereas the Schwartz criterion is minimized for 1 lag. We continue the analysis with  $k = 2$  and apply some tests of misspecification to the selected model. An LM test for first and fourth order autocorrelations, as well as a test for multivariate normality in the residuals are applied and showed in panel B in table 1. Details regarding the tests are found in Hansen and Juselius (1995). The autocorrelation tests reveal the presence of first order autocorrelation in the residuals with  $k = 2$ . Furthermore, the hypothesis of multivariate normality among the residuals is rejected. To reduce the autocorrelation, we add more lags to the model and the result of the misspecification tests in these cases are also shown in table 1. A model with 4 lags does not reveal signs of autocorrelation. However, multivariate normality among the residuals is still rejected.

One reason why the residuals do not appear to be normally distributed can be the presence of too many outliers among them, and an analysis of the estimated residuals from equation (5) reveals some large residuals. The largest one is present in the equation for

Table 1: Statistical results of the vector error correction model

| <b>Panel A: Information criteria</b> |              |          |
|--------------------------------------|--------------|----------|
| Lags                                 | Hannan-Quinn | Schwartz |
| 1                                    | -61.15       | -60.75   |
| 2                                    | -61.36       | -60.65   |
| 3                                    | -61.14       | -60.12   |
| 4                                    | -61.10       | -59.77   |

| <b>Panel B: Misspecification tests</b> |        |              |        |         |      |
|--|--------|--------------|--------|---------|------|
| 2 lags                                 |        |              |        |         |      |
| Autocorrelation                        | $LM_1$ | $\chi^2(36)$ | 56.68  | p-value | 0.02 |
|  | $LM_4$ | $\chi^2(36)$ | 45.36  | p-value | 0.14 |
| Normality                              |        | $\chi^2(12)$ | 264.68 | p-value | 0.00 |
| 3 lags                                 |        |              |        |         |      |
| Autocorrelation                        | $LM_1$ | $\chi^2(36)$ | 88.30  | p-value | 0.00 |
|  | $LM_4$ | $\chi^2(36)$ | 39.16  | p-value | 0.33 |
| Normality                              |        | $\chi^2(12)$ | 262.65 | p-value | 0.00 |
| 4 lags                                 |        |              |        |         |      |
| Autocorrelation                        | $LM_1$ | $\chi^2(36)$ | 29.25  | p-value | 0.78 |
|  | $LM_4$ | $\chi^2(36)$ | 34.06  | p-value | 0.54 |
| Normality                              |        | $\chi^2(12)$ | 248.27 | p-value | 0.00 |

*Note:* The misspecification tests shown in the table are tests for autocorrelation and multivariate normality, all reported by the software CATS.  $LM_1$  and  $LM_4$  are autocorrelation tests for first and fourth order autocorrelation, respectively. Details regarding the normality test are found in Hansen and Juselius (1995).

import prices for the observation of October 1982. Since the Swedish krona was devalued in this month, this is probably the cause of the large residual, and we account for this by including a dummy variable for this month in our model.<sup>7</sup> The inclusion of the dummy variable does not alter the results of the misspecification tests in table 1, i.e. there is some first order autocorrelation with 2 or 3 lags, but no autocorrelation with 4 lags.<sup>8</sup>

### 4.3 Roots of the characteristic polynomial

Another indication of I(2) symptoms in a cointegrated model is the presence of roots of the characteristic polynomial of the VAR model that are close to 1. The number of unit roots of the characteristic polynomial is the mirror image of the cointegration rank. With the rank equal to  $r$ , there are  $p - r$  unit roots and  $p - r$  common stochastic trends driving the system. If we have roots close to 1 for every choice of rank, this indicates the presence of common stochastic trends, possibly I(2) trends, in addition to the I(1) trends. In table 2, the largest roots of the characteristic polynomial for all choices of rank are shown. For a given rank,  $r$ ,  $p - r$  roots are equal to 1 by definition, but the  $p - r + 1^{th}$  root is very large for every rank, ranging from 0.92 to 0.99. This suggests the presence of an extra stochastic trend in the model, indicating that one or more of the variables have a double stochastic trend.

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<sup>7</sup>The standardized residual for the observation 1982:10 is 6.51 in the model with 4 lags. In November 1992, Sweden changed exchange rate policy and moved from a fixed to a floating exchange rate. In connection with this shift, the Swedish krona depreciated. In our estimations, this shift does not give rise to any outliers among the estimated residuals and therefore this observation is not treated with a dummy variable.

<sup>8</sup>We have also added dummy variables to handle other large residuals in the model. The non-normality of the residuals does not disappear when more dummy variables are included. In addition, the main conclusions from the analysis remain the same in the models including more dummy variables. We therefore proceed without adding any dummy variables other than the one accounting for the devaluation of October 1982.

Table 2: Roots of the characteristic polynomial

| Rank    |      |      |      |      |      |      |
|---------|------|------|------|------|------|------|
| $r = 1$ | 1.0  | 1.0  | 1.0  | 1.0  | 1.0  | 0.92 |
| $r = 2$ | 1.0  | 1.0  | 1.0  | 1.0  | 0.96 |      |
| $r = 3$ | 1.0  | 1.0  | 1.0  | 0.96 |      |      |
| $r = 4$ | 1.0  | 1.0  | 0.96 |      |      |      |
| $r = 5$ | 1.0  | 0.99 |      |      |      |      |
| $r = 6$ | 0.99 |      |      |      |      |      |

*Note:* The table shows the result for the model with 4 lags and one dummy variable for 1982:10. The result for the models with 2 or 3 lags are the same as in the table.

## 4.4 Long-run homogeneity between nominal wages and prices

Having established the probable presence of I(2) behaviour in the data, the variables causing this behaviour must be identified. In this section, we formally test the assumption made in section 3, that nominal wages are homogenous of degree zero in producer prices and consumer prices.

### 4.4.1 Determination of the cointegration rank

To determine the cointegration rank, we perform the Johansen (1988) trace test. The null hypothesis of the test is that the cointegration rank is equal to  $r$ , which is tested against the alternative hypothesis, that the rank is equal to  $p$ . The test statistic is

$$-2 \ln Q = -T \sum_{i=r+1}^p \ln(1 - \hat{\lambda}_i) \quad (6)$$

where  $\hat{\lambda}_i$  are the estimated eigenvalues of the matrix  $\Pi = \alpha\beta'$  and  $T$  is the number of observations. The result of the trace test is shown in table 3. The critical values at the 5 percent significance level and the p-values are those reported by the software CATS. The inclusion of dummy variables in the model may influence the distribution of the trace test and require other critical values, a question studied by Doornik et al. (1998). With dummy variables included in the estimations, the trace test is rejected too often and, therefore, the critical values are not large enough. In our case, we have only one dummy variable, taking

Table 3: Johansen's trace test of cointegration rank

|                     | $r = 0$ | $r \leq 1$ | $r \leq 2$ | $r \leq 3$ | $r \leq 4$ | $r \leq 5$ |
|---------------------|---------|------------|------------|------------|------------|------------|
| Eigenvalue          | 0.28    | 0.16       | 0.12       | 0.07       | 0.04       | 0.03       |
| Test statistic      | 190.92  | 108.24     | 65.98      | 33.16      | 16.35      | 6.55       |
| Critical value (5%) | 114.96  | 86.96      | 62.61      | 42.20      | 25.47      | 12.39      |
| p-value             | 0.00    | 0.00       | 0.03       | 0.33       | 0.47       | 0.46       |

*Note:* The table shows the result of Johansen's trace test for cointegration rank. Critical values and p-values are taken from the software CATS. The results presented are those for the model with 4 lags. The results for the models with 2 and 3 lags are the same.

the value one in one period and the value zero otherwise. Simulation results by Doornik et al. (1998) indicate that the effect of an unrestricted dummy variable on the distribution of the trace test is not generally large, yet this problem has to be kept in mind when we analyse the result of the trace test.

Using critical values at the 5 percent significance level, the hypotheses that the cointegration rank is  $r = 0$ ,  $r \leq 1$  and  $r \leq 2$  are all rejected, while the hypothesis that the rank is  $r \leq 3$  is not. The p-value for the hypothesis that  $r \leq 2$  is 0.03. The inclusion of the dummy variable implies that larger critical values should be used, which may have implications for the rejection of the hypothesis that  $r \leq 2$ . In table 3, we also show the estimated eigenvalues of the  $\Pi$ -matrix in the table. The two largest eigenvalues are 0.28 and 0.16 respectively, and the third is 0.12 and, thus, no large drop in the eigenvalues is observed that could indicate which of the eigenvalues are zero and which are not.<sup>9</sup>

In addition to the trace test, we analyse the recursively calculated trace test, which is shown in figure 2.<sup>10</sup> As seen in equation (6), the trace test statistic grows with the number

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<sup>9</sup>We have also applied the trace test to the model without a dummy variable. A Bartlett corrected trace test leads to the conclusion that the rank is 2. This result is robust for different lag lengths. Without the Bartlett correction, the rank is equal to 3 for the model with 3 and 4 lags.

<sup>10</sup>In figure 2, two graphs are shown. The upper one corresponds to the general model in equation (5), and the lower one to the so called concentrated model. In the vector error correction model in equation (5),  $\Delta X_t$  is regressed on  $X_t$  and  $\Delta X_{t-1}$ , where  $\Pi X_t$  defines the long-run relations in the model and  $\Gamma_1 \Delta X_{t-1}$  the short-run effects. In addition, the model is influenced by various deterministic components. In some cases, for example when deriving maximum likelihood estimators of a cointegrated VAR model, instead of

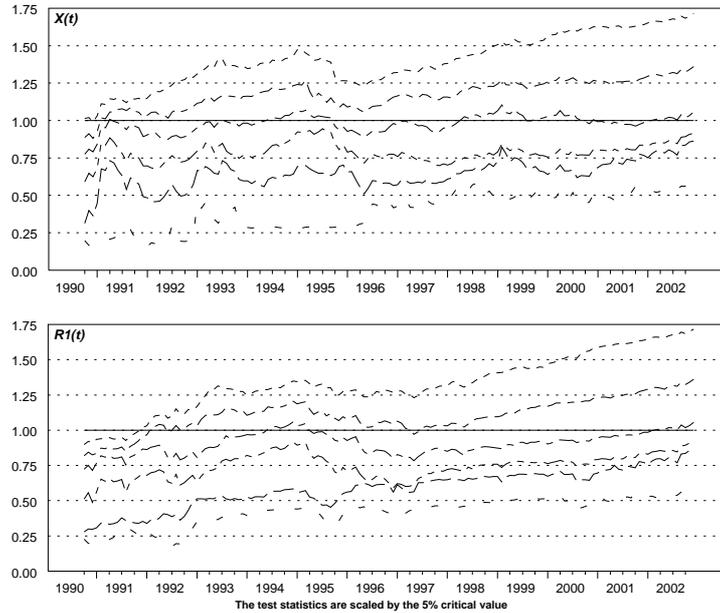


Figure 2: The recursively calculated trace test

of observations for eigenvalues larger than zero, while it is constant for eigenvalues equal to zero. In figure 2, the test statistics corresponding to the two largest eigenvalues show a clear upward sloping trend as the number of observations grows. The third eigenvalue grows initially, but then remains fairly constant when more observations are added to the estimations. In the lower graph, for the concentrated model, the third eigenvalue is below the critical value in a large segment of observations. In the end, the eigenvalue grows and exceeds the 5 percent critical value. Overall, the analysis points at a cointegration rank of 3, but due to the inclusion of the dummy variable, the case of a rank of 2 also warrants consideration.

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using the general model in (5), we concentrate out the short-run effects and the deterministic components and analyse this concentrated model instead. The concentrated model is often more stable than the general model, since the variation present in the lagged short-run terms is averaged out. The principles behind concentrating out the short-run effects and deterministic components are presented in e.g. Hendry and Juselius (2001). The derivations are explained in detail in Johansen (1995b).

Table 4: Test of long-run homogeneity

| 2 lags  |         |         |
|---------|---------|---------|
| Rank    | $r = 2$ | $r = 3$ |
| LR      | 1.18    | 6.87    |
| p-value | 0.55    | 0.08    |
| 3 lags  |         |         |
| Rank    | $r = 2$ | $r = 3$ |
| LR      | 1.96    | 10.38   |
| p-value | 0.38    | 0.02    |
| 4 lags  |         |         |
| Rank    | $r = 2$ | $r = 3$ |
| LR      | 16.59   | 19.66   |
| p-value | 0.00    | 0.00    |

*Note:* The table shows the result of the LR tests proposed in Johansen (1991) and Johansen and Juselius (1990, 1992). The test is  $\chi^2$  distributed with degrees of freedom equal to the restrictions placed on  $\beta$ .

#### 4.4.2 Test of long-run homogeneity

The assumption of long-run homogeneity between nominal wages and prices is tested with a likelihood ratio test (LR test) proposed in Johansen (1991) and Johansen and Juselius (1990, 1992). The null hypothesis of the test is formulated as

$$\mathcal{H} : \beta = H\varphi \tag{7}$$

where the same linear restrictions are imposed on the  $\beta$ -parameters in all  $r$  cointegration vectors.  $H$  is a design matrix with dimension  $(p \times s)$  and  $\varphi$  is an  $(s \times r)$ -matrix containing free parameters.  $p$ ,  $r$  and  $s$  are the number of variables, the cointegration rank and the number of free parameters, respectively. The LR test is asymptotically  $\chi^2$  distributed with degrees of freedom equal to the number of restrictions placed on the cointegration vectors.

The result of the homogeneity test is presented in table 4. The test is applied under the assumptions of  $r = 2$  and  $r = 3$  as well as in models with different lag length. In the model with 4 lags, we reject the null hypothesis of long-run homogeneity for  $r = 2$  and  $r = 3$ . However, this is not the case in the models with 2 and 3 lags. In these cases, we do not reject the hypothesis if  $r = 2$ , while we do reject it at the 5 percent significance level when

$r = 3$  and the lag length is 3. The support for the hypothesis of long-run homogeneity between wages and prices is thus not conclusive. A transformation to a model with real variables can be made even if the long-run homogeneity restriction is not valid, although there is a risk of losing some information in the model. Since a transformation to real terms simplifies the analysis, we apply a nominal-to-real transformation to the data and continue the analysis in an I(1) model.

## 5 Cointegration analysis of the I(1) model

### 5.1 The nominal-to-real transformation

The transformation to the I(1) model is made by deflating nominal wages, producer prices and consumer prices with one of the nominal variables. We chose to deflate with the producer price index, which generates a new data vector as

$$X_t = \left[ w_t^{real} \quad \Delta p_t^y \quad y_t \quad u_t \quad p_t^{c/y} \quad \tilde{p}_t^* \right]'. \quad (8)$$

Real wages,  $w_t^{real}$ , are defined as nominal wages less producer prices and the variable  $p_t^{c/y} = p_t^c - p_t^y$ . In the data vector, we need to include the growth rate in one of the nominal variables as a variable that captures the I(2)trend and, since the nominal series are deflated with producer prices, our choice is to include producer price inflation,  $\Delta p_t^y$ .<sup>11</sup> To obtain a model where no nominal variables appear in levels, we include the real exchange rate,  $\tilde{p}_t^*$ , instead of the import prices in the analysis of the I(1) model.<sup>12</sup>

The transformation of the original data vector to the vector of I(1) variables has important implications for the equations we want to estimate. Since the long-run homogeneity restriction is already imposed on the real wage equation in equation (2), this equation

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<sup>11</sup>The nominal-to-real transformation can be made by deflating the nominal series with any of the I(2) variables. The interpretation of some results alters, however, when deflating with different variables. The results and main conclusions from the analysis are unchanged if we deflate wages and prices with the consumer price index. If we deflate with nominal wages, our I(1) model will consist of a different set of variables, which alters the interpretation of the results. The real wage equation in (2) cannot be estimated in this setting since the specification does not include the variable  $p^{c/y}$ .

<sup>12</sup>Data on real exchange rate has been collected from OECD, Main Economic Indicators.

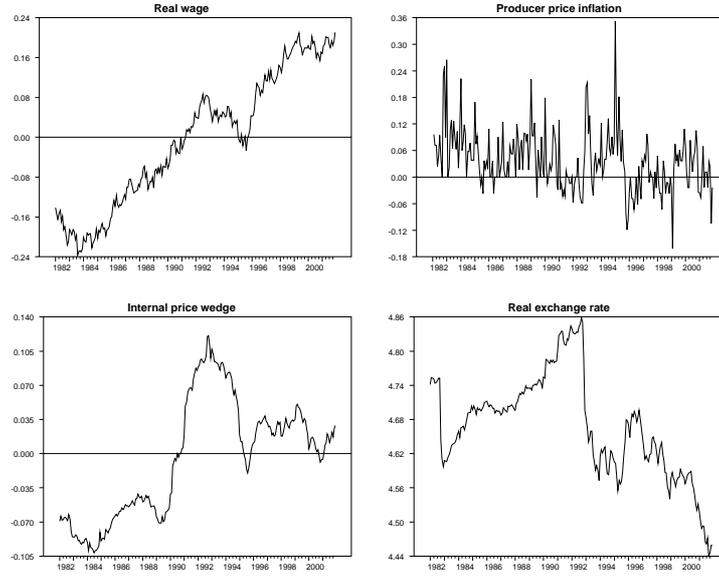


Figure 3: Graphs of real wages, producer price inflation, the internal price wedge and the real exchange rate

can be estimated in the transformed system as well. The price equation in equation (3), however, cannot be estimated in the I(1) framework since producer prices are no longer included in the model. Instead, we estimate a dynamic adjustment relation including the producer price inflation. The estimation of this relation is discussed in section 5.4.

The new variables in the I(1) model, real wages, producer price inflation, the price wedge and the real exchange rate, are presented in figure 3. Real producer wages exhibit an upward trend over the period, apart from a period in the beginning of the 1990s when the real wage falls, which reflects lower real wage costs for producers. The inflation rate in producer prices is presented in the top right graph. The monthly inflation rates have been multiplied by 12 to express the inflation rate on a yearly basis. A fall in the average inflation rate can be seen from 1995 onwards. In the lower left graph, the internal price wedge  $p_t^{c/y} = p_t^c - p_t^y$  is shown. The consumer price index measures prices of goods for final sale and the producer price index measures prices charged by manufacturers in the first link of the distribution chain. A rise in the variable  $p_t^{c/y}$  therefore indicates rising mark-ups for wholesalers, or a rising degree of monopoly in the distribution line. There

may, however, be other sources of variation in the variable throughout this period.

Inspecting  $p_t^{c/y}$  in figure 3, a striking feature is the rise in the series from the late 1980s to 1994. This increase in the price wedge probably arose from two sources. First, from figure 3, it can be seen that in the specific time period, producer prices fell. The manufacturing sector, in which the producer prices are measured, is one of the first sectors to suffer from recessionary tendencies affecting the economy. Initial actions towards such tendencies are to cut the firms' costs, which results in a falling producer price index and to lay off workers, which also lowers the costs and the price index. Combined with a low willingness to raise prices because of weaker demand in the export sector due to the international recession, the producer prices are expected to fall. In the same period, high domestic demand and an expansion in financial-sector lending led to high consumer price inflation in Sweden. Consequently, the variable  $p_t^{c/y}$  rose. Finally, real exchange rate is shown in the bottom right graph. The devaluation in 1982 and the depreciation in 1992 are clearly seen in the series.

## 5.2 The cointegration model

The basis of the analysis of the I(1) model is the same statistical model as for the nominal model in section 4.2. The vector error correcting model

$$\Delta X_t = \alpha\beta'X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \alpha\delta_0 t + \epsilon_t \quad (9)$$

is estimated based on the variable vector in equation (8). The disturbance term  $\epsilon_t$  is assumed to be independently and identically distributed as  $N(0, \Omega)$  and the model includes a deterministic trend, restricted to the cointegration relation. The two information criteria, the Hannan-Quinn and the Schwartz criteria, give inconclusive results to the question of the appropriate lag length of the model, as seen in table 5. The Hannan-Quinn criterion is minimized for 2 lags, while the Schwartz criterion is minimized for 1 lag. We have applied the test for autocorrelation and normality in the residuals to the model with both 1 and 2 lags but, in both cases, the model reveals problems with autocorrelation and non-normality of the residuals. Adding more lags to the model does not help to reduce or remove the

autocorrelation and we have to add a few dummy variables to correct for outliers among the observations. In all, we include three dummy variables in the model. The first one is for October 1982, when Sweden devaluated the Swedish krona. In addition to this, we include a variable for December 1992. In November the same year, Sweden moved from a fixed to a floating exchange rate and, in connection with this, the krona depreciated. In our data, however, a large residual appears for December, although the origin is the depreciation. Finally, a dummy variable for January 1995 is included.<sup>13</sup> This dummy variable considerably helps to reduce the autocorrelation in the model.<sup>14</sup> The final model for the analysis is then

$$\Delta X_t = \alpha\beta'X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \alpha\delta_0 t + \phi_1 d_{1982:10} + \phi_2 d_{1995:01} + \phi_3 d_{1992:12} + \epsilon_t. \quad (10)$$

The lag length is set to 3, i.e.  $k = 3$ . The result of misspecification tests regarding autocorrelation and normality in the estimated residuals of this model is presented in panel B in table 5. The residuals seem to be well behaved regarding autocorrelation properties, but the hypothesis of multivariate normality in the residuals is rejected. The rejection may arise due to large outliers. We have tried model specifications with several dummy variables included to account for this, but the test for normality is always rejected.<sup>15</sup>

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<sup>13</sup>The large residual appears in the equation for producer price inflation. In January 1995, Statistics Sweden introduced a new product classification system in calculating the producer price index. The new standard differs from the earlier considerably, which is the plausible cause of the large residual.

<sup>14</sup>Two large standardized residuals appear in the equation for real exchange rate in 1982:10 and 1992:12, with values of -6.89 and -5.22, respectively. In 1995:01 a standardized residual of 5.57 appears in the equation for producer price inflation.

<sup>15</sup>One possible source of non-normality in the residuals could be structural breaks in the model. We have tried model specifications where we have included a shift dummy variable in 1993. Such an inclusion does not help to remove the non-normality in the residuals. Further, it worsens the statistical properties of the model in terms of autocorrelation. A recursive analysis of the model with a shift dummy is also made. This model is not more stable in terms of constancy of the parameters than the model without a shift dummy variable.

Table 5: Statistical results of the vector error correction model with I(1) variables

| <b>Panel A: Information criteria 3 lags</b> |              |          |
|---|--------------|----------|
| Lags  | Hannan-Quinn | Schwartz |
| 1   | -55.25       | -54.84   |
| 2   | -55.36       | -54.64   |
| 3   | -55.14       | -54.13   |
| 4   | -54.95       | -53.63   |

| <b>Panel B: Misspecification tests</b> |        |              |         |         |      |
|--|--------|--------------|---------|---------|------|
| 3 lags and 3 dummy variables           |        |              |         |         |      |
| Autocorrelation                        | $LM_1$ | $\chi^2(36)$ | 46.47   | p-value | 0.11 |
|  | $LM_4$ | $\chi^2(36)$ | 35.35   | p-value | 0.50 |
| Normality                              |        | $\chi^2(12)$ | 165.078 | p-value | 0.00 |

*Note:* The misspecification tests shown in the table are tests for autocorrelation and multivariate normality, all reported by the software CATS.  $LM_1$  and  $LM_4$  are autocorrelation tests for first and fourth order autocorrelation, respectively. Details regarding the normality test are found in Hansen and Juselius (1995).

### 5.2.1 Determination of the cointegration rank

The determination of the cointegration rank is mainly by the application of the Johansen (1988) trace test, the result of which is shown in table 6. The hypotheses that  $r = 0$  and  $r \leq 1$  are rejected with p-values equal to 0.00. The hypothesis that  $r \leq 2$  has a p-value of 0.11 and is not rejected at the 10 percent significance level. Examining the eigenvalues, we see that there is a large drop from the second to third eigenvalues, from 0.19 to 0.08, which is also an indication that two of the eigenvalues are different from zero.<sup>16</sup>

The recursively calculated trace test is shown in figure 4, where the upper and lower graphs are calculations for the general model and the model concentrated for short-run effects, respectively. The two largest eigenvalues show a clear upward trend when more observations are added to the estimations. For the concentrated model, the third eigenvalue seems to be constant, and insignificant, throughout the period. As well as the trace test,

<sup>16</sup>A Bartlett corrected trace test in a model without any dummy variables has been conducted. The test clearly indicates that the rank is 2. The result is robust for different lag lengths and deterministic components.

Table 6: Johansen's trace test of cointegration rank in the I(1) model

|                     | $r = 0$ | $r \leq 1$ | $r \leq 2$ | $r \leq 3$ | $r \leq 4$ | $r \leq 5$ |
|---------------------|---------|------------|------------|------------|------------|------------|
| Eigenvalue          | 0.31    | 0.28       | 0.08       | 0.07       | 0.05       | 0.03       |
| Test statistic      | 233.42  | 141.15     | 59.67      | 38.27      | 21.51      | 8.09       |
| Critical value (5%) | 114.95  | 86.96      | 62.61      | 42.20      | 25.46      | 12.38      |
| p-value             | 0.00    | 0.00       | 0.11       | 0.14       | 0.16       | 0.30       |

*Note:* The table shows the result of Johansen's trace test for cointegration rank. Critical values and p-values are taken from the software CATS.

the recursively calculated trace test indicates that the appropriate cointegration rank is 2.

### 5.3 Stationarity and weak exogeneity

In this section, we test different hypotheses on the parameters in the matrix  $\Pi = \alpha\beta'$ . LR tests for testing hypotheses on  $\alpha$  and  $\beta$  are derived in Johansen (1991) and Johansen and Juselius (1990, 1992). All tests are  $\chi^2$  distributed with degrees of freedom equal to the number of restrictions placed on the parameters in the cointegration vector  $\beta$  or the speed of adjustment terms in  $\alpha$ .

Based on graphical considerations of the variables, a restricted linear trend is included in the estimations of the cointegrated VAR model. We test the validity of this inclusion by considering the parameters in the long-run relations,  $\beta$ . A hypothesis of excluding the time trend is formulated as  $\mathcal{H} : \beta = H\varphi$  and the test yields a test statistic of 24.92 and a p-value of 0.00. We reject the null hypothesis of exclusion of the trend and leave the trend in the model specification.

All variables are also tested for stationarity with a null hypothesis as

$$\mathcal{H} : \beta = (H_1\varphi_1, \varphi_2). \quad (11)$$

With this hypothesis, we test if either of the two cointegration vectors consists of solely one stationary variable.  $H_1\varphi_1$  defines the hypothesis of a stationary variable in the first cointegration vector and the parameters in the other relation,  $\varphi_2$ , are left unrestricted. The

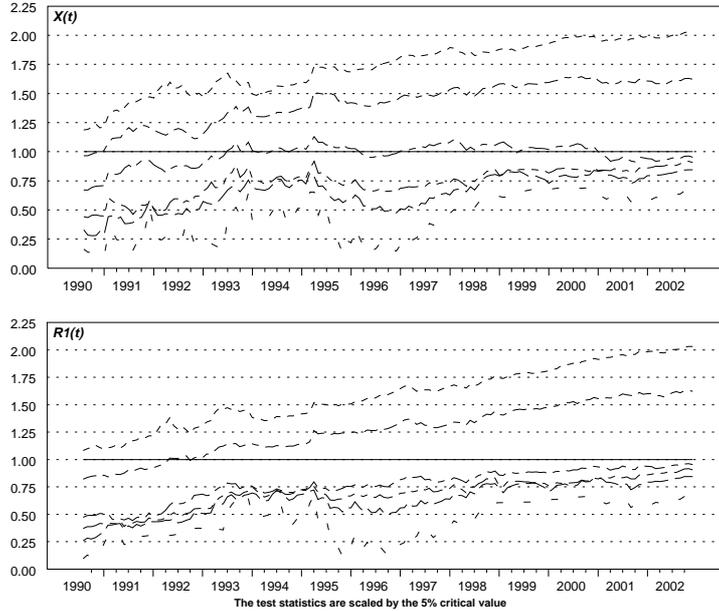


Figure 4: The recursively calculated trace test for the model with I(1) variables

result of the stationarity tests is presented in panel A in table 7. None of the variables are stationary, since the test is rejected in all cases.

A long-run weakly exogenous variable influences the development of other variables in the model, but is not influenced by any other variable itself. Weak exogeneity can be tested by testing hypotheses of the adjustment parameters,  $\alpha$ , in  $\Pi = \alpha\beta'$ . If an  $\alpha$ -parameter corresponding to a certain variable and cointegration relation is equal to zero, that variable does not adjust to developments in that cointegration relation. If the  $\alpha$ -parameters of a certain variable corresponding to all cointegration relations are zero, the variable does not adjust to developments in any long-run relation, which means that it is weakly exogenous. The null hypothesis of weak exogeneity can be formulated as

$$\mathcal{H} : R'\alpha = 0 \tag{12}$$

where  $R$  is a  $(s \times p)$ -matrix extracting the elements in  $\alpha$  equal to zero,  $s$  is the number of restrictions placed on  $\alpha$  and  $p$  the number of variables.

Test results for weak exogeneity are presented in panel B in table 7. Weak exogeneity

Table 7: Hypothesis test on  $\alpha$  and  $\beta$ 

| <b>Panel A: Stationarity</b> |            |              |       |       |         |               |
|------------------------------|------------|--------------|-------|-------|---------|---------------|
|                              | $w^{real}$ | $\Delta p^y$ | $y$   | $u$   | $p^c/y$ | $\tilde{p}^*$ |
| LR                           | 76.90      | 30.49        | 72.42 | 71.15 | 72.09   | 68.08         |
| p-value                      | 0.00       | 0.00         | 0.00  | 0.00  | 0.00    | 0.00          |

| <b>Panel B: Weak exogeneity</b> |            |              |      |       |         |               |
|---------------------------------|------------|--------------|------|-------|---------|---------------|
|                                 | $w^{real}$ | $\Delta p^y$ | $y$  | $u$   | $p^c/y$ | $\tilde{p}^*$ |
| LR                              | 9.02       | 27.17        | 6.32 | 44.34 | 11.52   | 0.87          |
| p-value                         | 0.01       | 0.00         | 0.04 | 0.00  | 0.00    | 0.65          |

*Note:* The table shows the result of the LR tests proposed in Johansen (1991) and Johansen and Juselius (1990, 1992). The tests are  $\chi^2$  distributed with degrees of freedom equal to the restrictions placed on  $\alpha$  and  $\beta$ .

is rejected at the 5 percent significance level for all variables except the real exchange rate, where the p-value is 0.65. The p-value of the test of weak exogeneity of productivity is 0.04, hence it is a borderline case of rejection at the 5 percent significance level. That these two variables are weakly exogenous is in accordance with economic theory. Import prices, that were originally included in the nominal model, should be weakly exogenous in a small open country like Sweden, since import prices influence Swedish economic variables, whereas the development in the Swedish economy does not influence prices in the rest of the world. For the real exchange rate, the weak exogeneity is not as obvious since the real exchange rate is also influenced by the domestic price level. However, it is an economically plausible result, that the real exchange rate is weakly exogenous. Productivity is also a variable that is thought to be weakly exogenous in the long-run, since productivity, or technological development, is the driving force behind long-run growth. Weak exogeneity is however rejected in table 7. This result may emanate from rigidities in the labour market.

## 5.4 Identification and estimation of the long-run relations

The purpose of this section is to impose restrictions on the two cointegration relation in order to identify the long-run structure of the model. Over-identifying restrictions

Table 8: Hypotheses on the long-run parameters

| <b>Panel A: 3 lags</b>          |            |              |       |      |           |               |         |      |         |
|---------------------------------|------------|--------------|-------|------|-----------|---------------|---------|------|---------|
|                                 | $w^{real}$ | $\Delta p^y$ | $y$   | $u$  | $p^{c/y}$ | $\tilde{p}^*$ | $trend$ | LR   | p-value |
| $\mathcal{H}_1$                 | 1.00       |              | -2.48 | 0.76 | -0.99     |               | 0.003   | 2.82 | 0.09    |
| $\mathcal{H}_2$                 | 0.14       | 1.00         |       |      |           | -0.0001       | -0.0001 | 3.23 | 0.20    |
| $\mathcal{H}_3$                 |            | 1.00         |       | 0.30 |           | 0.01          | 0.0003  | 1.47 | 0.48    |
| $\mathcal{H}_1 + \mathcal{H}_2$ |            |              |       |      |           |               |         | 6.02 | 0.11    |
| $\mathcal{H}_1 + \mathcal{H}_3$ |            |              |       |      |           |               |         | 4.33 | 0.23    |

| <b>Panel B: 4 lags</b>          |            |              |       |      |           |               |         |      |         |
|---------------------------------|------------|--------------|-------|------|-----------|---------------|---------|------|---------|
|                                 | $w^{real}$ | $\Delta p^y$ | $y$   | $u$  | $p^{c/y}$ | $\tilde{p}^*$ | $trend$ | LR   | p-value |
| $\mathcal{H}_1$                 | 1.00       |              | -1.93 | 0.76 | -1.07     |               | 0.002   | 1.11 | 0.29    |
| $\mathcal{H}_2$                 | 0.39       | 1.00         |       |      |           | -0.07         | -0.0005 | 6.70 | 0.04    |
| $\mathcal{H}_3$                 |            | 1.00         |       | 0.60 |           | -0.03         | 0.0001  | 3.68 | 0.16    |
| $\mathcal{H}_1 + \mathcal{H}_2$ |            |              |       |      |           |               |         | 7.81 | 0.05    |
| $\mathcal{H}_1 + \mathcal{H}_3$ |            |              |       |      |           |               |         | 5.29 | 0.15    |

on the parameters in the cointegration vectors are tested using the LR test proposed by Johansen and Juselius (1992). We also simultaneously test weak exogeneity restrictions on the adjustment parameters,  $\alpha$ , in order to specify a suitable model. The over-identifying restrictions are derived from the equations in the theoretical model in section 3.

It is straightforward to formulate a hypothesis based on the real wage equation in equation (2). The hypothesis is formulated as  $\mathcal{H}_1 : \beta = (H_1\varphi_1, \varphi_2)$  and the result of the test is found in the first row in panel A of table 8. The p-value of  $\mathcal{H}_1$  is 0.09 and the null hypothesis is not rejected at the 5 percent significance level.

Turning to the second equation, equation (3) cannot directly be estimated since producer prices are not included in levels in the I(1) model. Instead, we estimate a dynamic adjustment relation, as described in Juselius (2003a), for the development of producer price inflation,  $\Delta p^y$ . In the theoretical relations,  $p^y$  is specified as  $p_t^y = \theta_0 + \theta_1 w_t^{nom} + (1 - \theta_1)p_t^*$  and therefore, we formulate a hypothesis,  $\mathcal{H}_2$  in table 8, where producer price inflation adjusts to changes in real wages and the real exchange rate. The p-value of the hypothesis is 0.20 and the hypothesis is not rejected.

We also try an alternative formulation of the second equation. With wage rigidity

present in the labour market, real wages may not adjust to changes in price inflation. Instead, wage rigidity may lead to disequilibria in the labour market, appearing as changes in the unemployment rate. The hypothesis  $\mathcal{H}_3$  in table 8 is formulated along this line, where price inflation is adjusting to import prices, as before, and the unemployment level instead of real wage. Testing this hypothesis yields a p-value of 0.48, and the hypothesis is not rejected.

To check the robustness of the statistical results, we have made the analysis for other model formulations in terms of lag lengths and deterministic components. So far, the results are not sensitive to these choices. To analyse the robustness regarding the over-identifying restrictions on the cointegration vectors, we show the test results when including an additional lag. This model has the same statistical properties with respect to autocorrelation and normality. The test results regarding hypotheses on  $\beta$  for the model with 4 lags are shown in panel B in table 8. With 4 lags, the hypotheses  $\mathcal{H}_1$  and  $\mathcal{H}_3$  are not rejected, at the 5 percent significance level. However,  $\mathcal{H}_2$  is rejected with a p-value of 0.04. Since the test of  $\mathcal{H}_2$  has a lower p-value than the test of  $\mathcal{H}_3$  with 3 lags, and since  $\mathcal{H}_2$  is rejected in the model with 4 lags, we turn in favour of the hypothesis  $\mathcal{H}_3$ . Joint tests of the hypotheses are also reported in table 8. A joint test of  $\mathcal{H}_1$  and  $\mathcal{H}_3$  is not rejected in any of the models.<sup>17</sup>

$\mathcal{H}_1$  and  $\mathcal{H}_3$  together identify the two cointegration vectors and pass the rank condition for formal identification in Johansen (1995a). The two hypotheses also yield an inter-

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<sup>17</sup>The joint test of  $\mathcal{H}_1$  and  $\mathcal{H}_3$  is tested by formulating a hypothesis

$$\mathcal{H} : \beta = (H_1\varphi_1, H_2\varphi_2) \tag{13}$$

with design matrices  $H_1$  and  $H_2$  as

$$H_1 = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}, H_2 = \begin{pmatrix} 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{pmatrix}. \tag{14}$$

Table 9: Estimated cointegration vectors in a model with over-identifying restrictions

|            | $w^{real}$  | $\Delta p^y$ | $y$             | $u$            | $p^{c/y}$       | $\tilde{p}^*$  | $trend$            |
|------------|-------------|--------------|-----------------|----------------|-----------------|----------------|--------------------|
| $\beta'_1$ | 1.00<br>(-) |              | -2.52<br>(0.29) | 0.76<br>(0.27) | -1.00<br>(0.19) |                | 0.004<br>(0.001)   |
| $\alpha_1$ | -0.06**     |              | 0.04**          | 0.03**         |                 | (-)            |                    |
| $\beta'_2$ |             | 1.00<br>(-)  |                 | 0.28<br>(0.17) |                 | 0.01<br>(0.05) | 0.0003<br>(0.0001) |
| $\alpha_2$ |             | -0.51**      |                 |                | -0.03**         |                |                    |
| LR         | 3.90        |              |                 |                |                 |                |                    |
| p-value    | 0.27        |              |                 |                |                 |                |                    |

*Note:* The estimates in the table are from a model where the real exchange rate is weakly exogenous. Standard errors for the parameters in  $\beta$ , reported by the software CATS, are in parentheses. For the  $\alpha$ -parameters, \* indicates significance at the 5 percent level and \*\* indicates significance at the 1 percent level.

pretable matrix of adjustment parameters with error-correcting behaviour. In the presence of a weakly exogenous variable, fully efficient estimates of the long-run parameters in  $\beta$  are generated from estimations of a partial model conditioning on the weakly exogenous variable, as shown by Johansen (1992). After identifying the real exchange rate as weakly exogenous, the model is re-estimated conditioned on the real exchange rate and a further analysis is carried out in the partial model.<sup>18</sup> We have also carried out a recursive analysis of the identified, partial model to check the constancy of the parameters over the sample period. The recursive tests are described in Hansen and Juselius (1995). Generally, the concentrated model, corrected for short-run adjustment and deterministic components, behaves satisfactorily in terms of stability. The full model is to some extent more volatile.<sup>19</sup> In table 9, we present the estimates of the  $\alpha$  and  $\beta$ -parameters of the identified model.

The first relation,  $\beta'_1$ , specifies a relation between wages, labour productivity, unem-

<sup>18</sup>Conditional on the real exchange rate, we have tested if productivity is weakly exogenous as well. This test is rejected at the 5 percent significance level.

<sup>19</sup>When we statistically test the constancy of the eigenvalues, we do not reject the null hypothesis of constant eigenvalues for the first or the second eigenvalue, nor for the sum of them, in the concentrated model. In the full model, this test is borderline rejected for the first eigenvalue at 5 percent significance level but not at the 1 percent level. Recursively calculated  $\beta$ -vectors and log-likelihood functions are stable for the concentrated model, while the full model is more volatile over the sample period.

ployment and the price wedge normalized on real wages.  $\beta_1'$  is an estimate of equation (2) in section 3, and the parameters in the cointegration relations in table 9 can be compared to the theoretical parameters. The relation between real wages and productivity, captured in the parameter  $\gamma_2$ , is expected to be positive. The empirical estimates support this expectation, since the estimate of  $\gamma_2$  to  $-2.52$  corresponds to a positive relation between the variables. A rise in productivity leads to a rise in real wages, but the empirical relation is not one-to-one as economic theory predicts. The hypothesis of a one-to-one relation between the two variables is rejected. From the parameter estimates, a rise in productivity leads to a more than twice as high a rise in real wages. This implies that real wages have grown too fast compared to the growth in productivity, a feature widely believed to have been present in the Swedish economy during the 1980s. From the theoretical relation, we expect a negative relation between real wages and the unemployment level, which is also found in the empirical estimates.  $\gamma_3$ , which describes the impact of higher unemployment, is estimated to be  $0.76$ . The sign is in accordance with our expectations. Finally, the parameter in front of the internal price wedge,  $\gamma_1$ , is estimated to be  $-1.00$ .<sup>20</sup> Since we interpret  $\gamma_1$  as a measure of trade union power, the result implies strong union power as the real wage development has closely followed the development in consumer prices.

Turning to the adjustment parameters in  $\alpha_1$ , we see that real wages, productivity and unemployment adjust to a deviation in the long-run relation. Real wages and productivity are error-correcting, i.e. the variables adjust back towards equilibrium after a shock to the relation. Unemployment is not error-correcting, so developments in that variable do not work in the direction that restores equilibrium. The dynamics of unemployment are, however, in line with the developments in real wages and productivity. If there is a positive deviation in the long-run relationship due to a disturbance in real wages, unemployment responds by increasing and if there is a negative deviation due to a positive productivity disturbance, unemployment decreases. Moreover, the magnitude of the adjustment parameter to unemployment is smaller than the magnitude of the  $\alpha$ -parameters corresponding to the error-correcting variables so, in the long-run, equilibrium will be restored. The ad-

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<sup>20</sup>The parameter estimate is not significantly different from 1.

justment parameters imply that there are rigidities present in the Swedish labour market, since both productivity and unemployment react to developments in the real wage relation. With no rigidities, disequilibria in the labour market would lead to adjustments only in wages or in nominal variables.

The second relation,  $\beta'_2$ , describes the interrelation of price inflation, unemployment and import prices. Due to the presence of I(2) and the nominal-to-real transformation, the relation cannot be interpreted as one of the theoretical relations in section 3, since inflation is included instead of prices in levels. The relation specifies a negative relation between price inflation and unemployment, as in the traditional Phillips curve. A rise in unemployment puts less pressure on prices, and inflation decreases. The long-run relation between inflation and unemployment may suggest rigidities in the labour market. Producers then respond to various shocks affecting their business by adjusting the employment level instead of responding by adjusting wages as would be expected in a labour market without rigidities. Those rigidities, alongside the I(2) behaviour in the data, are the probable cause of the second cointegration relation not reflecting the pure price-setting behaviour that the theoretical relations predict. In  $\beta_2$ , the positive relation between producer price inflation and the real exchange rate implies that an increase in import prices results in increased producer price inflation, which is in line with the theoretical relations in section 3. The parameter is, however, not significantly different from zero.

Only inflation and the price wedge have significant  $\alpha$ -parameters and adjust to deviations in this relation. Producer price inflation is error-correcting and the adjustment in the variable is very large. Comparing the adjustment parameter for inflation of  $-0.51$  to the other adjustment parameters in the two relations, price inflation adjusts more than ten times faster than the other variables. The adjustment in the price wedge to deviations in this second relation probably originates from the close connection between the price wedge and producer price inflation, where  $p_t^{c/y} = p_t^c - p_t^y$ . Note that this second relation is a nominal relation, describing the dynamics in prices, and that the real variables, real wages, productivity and unemployment, do not adjust to deviations in this relation.

## 6 Common trends analysis

Instead of describing the cointegration model in terms of the long-run steady state relations and the adjustment processes towards these relations, we can describe the model in terms of the underlying common stochastic trends. By imposing identifying restrictions on parameters in the common trends model, the underlying trends can be given an economic interpretation as structural shocks.

In order to identify the common stochastic trends, we invert the autoregressive vector error correction model

$$\Delta X_t = \alpha\beta'X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \epsilon_t. \quad (15)$$

into the moving average model<sup>21</sup>

$$X_t = C \sum_{i=1}^t \epsilon_i + C^*(L)\epsilon_t + \tilde{X}_0. \quad (16)$$

This follows from the Granger representation theorem (Engle and Granger, 1987). In equation (16),  $X_t$  is determined by the stochastic trends that are built up by  $\sum_{i=1}^t \epsilon_i$ , the cumulative sum of innovations to the  $p$  variables.  $X_t$  is also determined by the stationary stochastic components in  $C^*(L)\epsilon_t$  and initial values in  $\tilde{X}_0$ . The parameter matrix  $C$  is related to the parameters in the vector error correction model as

$$C = \beta_{\perp}(\alpha'_{\perp}\Gamma\beta_{\perp})^{-1}\alpha'_{\perp}. \quad (17)$$

$\beta_{\perp}$  and  $\alpha_{\perp}$  are orthogonal components of  $\beta$  and  $\alpha$  and  $\Gamma = I - \Gamma_1 - \dots - \Gamma_{k-1}$ .  $C^*(L) = \sum_{i=0}^{\infty} C_i^* L^i$ , where  $L$  is the lag operator, and  $C_i^* = \alpha(\beta'\alpha)^{-1}(I_r + \beta'\alpha)^i\beta'$ .<sup>22</sup>

In a cointegrated model,  $C$  has a reduced rank of  $p - r$ , which implies that of the  $p$  innovations in  $\sum_{i=1}^t \epsilon_i$ , only  $p - r$  of them translate into stochastic trends with permanent effects on the variables in  $X_t$ . Consequently, with  $p$  variables and  $r$  cointegration relations, there are  $p - r$  common trends driving the variables. These common driving trends are defined by  $\alpha'_{\perp} \sum_{i=1}^t \epsilon_i$  and the remaining part of  $C$ , i.e.  $\beta_{\perp}(\alpha_{\perp}\Gamma\beta_{\perp})^{-1}$ , can be interpreted

<sup>21</sup>To simplify the illustration of the common trends model, we disregard the deterministic components.

<sup>22</sup>For details about the derivations of moving average representations for vector autoregressive models, see Johansen (1995b).

as loadings to the common trends, i.e. the coefficients with which the common trends influence the variables in  $X_t$ .

The moving average representation in equation (16) is a reduced form of a common trends model. It consists of the cumulative sum of the estimated residuals, but since these are correlated, it is hard to interpret them as structural shocks. The goal of the structural common trends analysis is then to recover the structural shocks from the estimated residuals. To do this, we assume a linear relation between the structural shocks,  $u_t$ , and the residuals,  $\epsilon_t$ , as

$$u_t = B\epsilon_t \quad (18)$$

and additionally, that  $E(u_t u_t') = I_p$ , so that the structural shocks are uncorrelated with unit variance. By multiplying equation (16) with  $B^{-1}B$  a structural version of the moving average representation is given as

$$\begin{aligned} X_t &= CB^{-1}B \sum_{i=1}^t \epsilon_i + C^*(L)B^{-1}B\epsilon_t + \tilde{X}_0 \\ &= CB^{-1} \sum_{i=1}^t u_i + C^*(L)B^{-1}u_t + \tilde{X}_0. \end{aligned} \quad (19)$$

The identification of the structural shocks rests on finding a suitable matrix  $B$ . We impose an identifying structure on the common trends model with the method proposed by Gonzalo and Ng (2001).<sup>23</sup> First, we decompose the structural shocks into permanent and transitory shocks as  $u_t = \begin{pmatrix} u_t^p & u_t^t \end{pmatrix}'$ .  $u_t^p$  then consists of  $p - r$  shocks with permanent effects on  $X_t$  and  $u_t^t$  consists of  $r$  shocks with transitory effects.<sup>24</sup> The structural version of the common trends model in equation (19) can be rewritten as

$$X_t = \Upsilon \sum_{i=1}^t u_i^p + C^*(L)B^{-1} \begin{pmatrix} u_t^p \\ u_t^t \end{pmatrix} + \tilde{X}_0. \quad (20)$$

if we define  $CB^{-1} = \begin{pmatrix} \Upsilon & 0_{p \times r} \end{pmatrix}$ .  $\Upsilon$  is then an impact matrix that contains the long-run effect on the variables in  $X_t$  after a shock to the stochastic trends with a permanent effect.

<sup>23</sup>For another identification method of common trends models, see Warne (1993).

<sup>24</sup>The decomposition is accomplished by multiplying the estimated residuals with a matrix  $G = \begin{bmatrix} \alpha'_\perp & \alpha'\Omega^{-1} \end{bmatrix}'$ , where  $\alpha'_\perp \epsilon_t$  are the permanent innovations and  $\alpha'\Omega^{-1}\epsilon_t$  are the transitory innovations.  $\Omega$  is the variance-covariance matrix for the estimated residuals.

By imposing restrictions on the long-run effect in  $\Upsilon$ , the structural model can be exactly identified. Based on these restrictions, the matrix  $B$  can be calculated.<sup>25</sup>

Introducing the matrix  $B$  gives  $p^2$  new parameters to be estimated. The assumption that  $u_t$  are uncorrelated with unit variance,  $\Sigma_u = B\Omega B' = I_p$ , restricts  $\frac{p+p(p-1)}{2}$  parameters. Moreover, the assumption that the transitory shocks have no long-run impact on the variables in  $X_t$  restricts  $(p-r)r$  additional parameters. However, we still need to impose  $\frac{(p-r)(p-r-1)}{2} + \frac{r(r-1)}{2}$  restrictions on  $\Upsilon$  to identify the structural model. Of these restrictions,  $\frac{(p-r)(p-r-1)}{2}$  must be added to identify the shocks with permanent effect on  $X_t$  and  $\frac{r(r-1)}{2}$  must be added to identify the transitory shocks.

## 6.1 Identification and interpretation of the structural shocks

In this section, we identify and analyse the common driving trends in the empirical model. The focus lies on the common trends that have permanent effects on the variables, and we identify these trends by placing zero-restrictions on the long-run impact matrix  $\Upsilon$ .<sup>26</sup> The model consists of six variables and two cointegration relations and, therefore, there are four common trends with permanent effects on  $X_t$ . In all, we have to impose 6 restrictions on  $\Upsilon$  to identify the shocks with permanent effects.<sup>27</sup>

The economy is continuously influenced by a variety of shocks, affecting the economy in various ways. A natural division of the shocks is between real shocks and nominal shocks, where real shocks, for example productivity shocks, are expected to have permanent effects whereas nominal shocks are expected to have temporary effects only. Apart from domestic real and nominal shocks, an open country like Sweden is likely to be influenced by foreign shocks, which can originate from either real or nominal sources, giving rise to different effects on the Swedish economy.

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<sup>25</sup>Technically, this can be accomplished by gathering the identifying restrictions in a matrix  $K$ , that is multiplied with the decomposed innovations. To ensure that the structural shocks are orthogonal, they are premultiplied with a Choleski decomposition of  $K^{-1}G\Omega G'K'^{-1}$ . The matrix  $B$  is calculated as  $B = M^{-1/2}KG$ , where  $M^{-1/2}$  is the matrix from the Choleski decomposition.

<sup>26</sup>The dimension of  $\Upsilon$  is  $(p \times (p-r))$ , where the columns correspond to the  $p-r$  common stochastic trends and the rows to the  $p$  variables.

<sup>27</sup>To identify the shocks with permanent effect on  $X_t$  we have to impose  $\frac{(p-r)(p-r-1)}{2} = 6$  restrictions on  $\Upsilon$ .

In the previous analysis, we found the real exchange rate to be weakly exogenous. This has an important implication for the common trends analysis since it implies that one of the driving trends consists of shocks to this variable. The weak exogeneity also implies that the remaining three common trends do not have a long-run impact on the real exchange rate. Three of the six zero-restrictions can thus be put in the row corresponding to the real exchange rate. The three additional restrictions that are needed do not follow directly from the original model. Based on economic theory, we expect that, in the long-run, productivity is determined by the overall technological level in the economy and affected only by real shocks. For that reason, we place two of the remaining three restrictions in the row for productivity. The last restriction is placed on unemployment, since, in the long-run, it is expected to be affected by structural factors in the labour market, and not by shocks to the remaining variables. These six restrictions yield a long-run impact matrix as

$$\Upsilon = \begin{pmatrix} * & * & * & * \\ * & * & * & * \\ * & * & 0 & 0 \\ * & * & * & 0 \\ * & * & * & * \\ * & 0 & 0 & 0 \end{pmatrix}. \quad (21)$$

The columns in  $\Upsilon$  correspond to the four stochastic trends and the rows correspond to the variables.<sup>28</sup> An asterisk (\*) indicates that the parameter in the long-run matrix is free to vary and a zero is a zero-restriction.<sup>29</sup>

A shock to the first trend is allowed to influence all of the variables in the long-run. The weak exogeneity of the real exchange rate implies that the first trend is interpreted to be shocks to the real exchange rate. The second trend is interpreted to be shocks to productivity, since all domestic variables, including productivity, are allowed to be permanently affected by the trend. Thus, the trend influences all variables but the real

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<sup>28</sup>The variables are ordered as  $[w_t^{real} \ \Delta p_t^y \ y_t \ u_t \ p^c p_t^y \ \tilde{p}^*]'$ .

<sup>29</sup>With the restrictions in  $\Upsilon$  in equation (21), we can calculate the identification matrix  $B$  with the method proposed by Gonzalo and Ng (2001). Details about the identification scheme can be found there.

exchange rate in the long-run. The third trend has zero-restrictions imposed on the real exchange rate and productivity, while unemployment is allowed to react in the long-run. We interpret this trend as consisting of shocks originating from the labour market. Finally, the fourth trend, with an additional restriction on unemployment, is harder to interpret. Turning to the variables in the model, import prices and productivity are assumed to be the source of the first and second trends and real wages and unemployment to be the source of the third trend. Shocks to inflation are not expected to give rise to any long-run effects in the economy, which leaves the price wedge as a plausible origin for the fourth trend. This variable represents the relative degree of monopoly between producers and wholesalers, and a shock to the variable would represent a change in this monopoly situation.

## 6.2 Impulse response analysis

This section contains the results of an impulse response analysis of the identified common trends model in the previous section. The two identifying cointegration relations found in section 5.4 as well as the weak exogeneity condition of the real exchange rate are imposed on the model before the impulse response functions are calculated. In figures 5 to 8, we present the impulse responses after shocks to the four common trends. Besides the impulse responses, all graphs show bootstrapped confidence bands with one standard deviation margin for the estimations.<sup>30</sup>

Figure 5 shows impulse response functions after a shock to the real exchange rate. This shock leads to a negative response of the real exchange rate itself, as seen in the graph for the real exchange rate. From the response of producer price inflation, we see that the effect on price inflation is temporarily positive. The rise in domestic prices leads to lower

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<sup>30</sup>The bootstrapped standard deviation bands are constructed in the following way. Parameter estimates and fitted residuals from the estimation of the vector error correction model in equation (10) are saved. By random sampling with replacement among these residuals, a new set of residuals is collected and a new set of data is calculated based on the initial parameter estimates. After that, the model is re-estimated and impulse responses are calculated and stored. For all re-estimations, the number of cointegration vectors and the restrictions on the  $\beta$ - and  $\alpha$ -parameters from the original model are fixed. The procedure is repeated 500 times and empirical standard deviations for the impulse response functions are calculated.

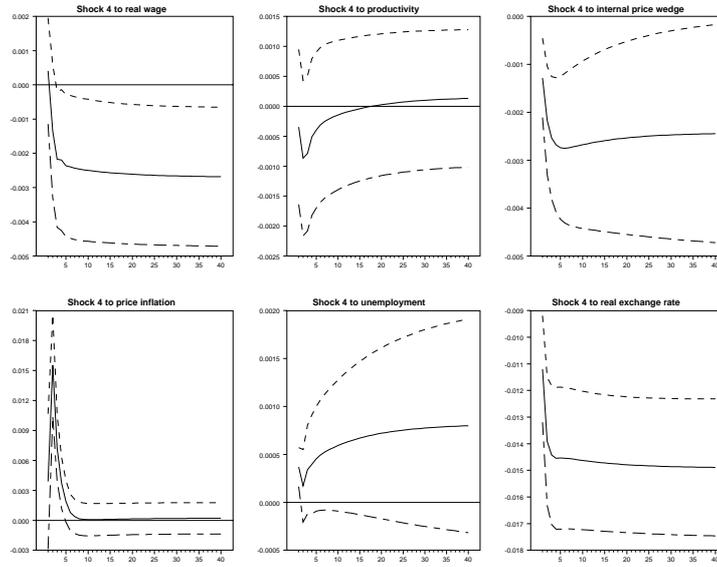


Figure 5: Impulse responses after a shock to the real exchange rate

real producer wages. This pattern of responses can be due to the fact that a rise in import prices is transmitted to the Swedish producer market and causes a temporary increase in producer inflation. This relation between import prices and producer prices is also seen in the cointegration relation describing producer price inflation in section 5.4. The effect on the price inflation lasts for about eight months, and after that, a new equilibrium level for real wages is reached. The internal price wedge reacts negatively, in line with the development of producer prices. The reactions of productivity and unemployment are not significant.

The second common trend is interpreted to be a domestic productivity shock. In figure 6, we see that productivity reacts positively, and significantly, to this shock. The effect on real wages is also positive, which is clearly in line with the interpretation of a productivity shock. The initial reaction of unemployment is positive, while the long-run effect is negative, but the confidence bands are large and only the initial response is significant. The effect of a productivity shock on producer price inflation is zero, which also confirms the interpretation of the shock. Finally, the long-run effect on the price wedge is not significant. The initial effect, however, is negative, which indicates that inflation in consumer prices is weakening after a positive productivity shock.

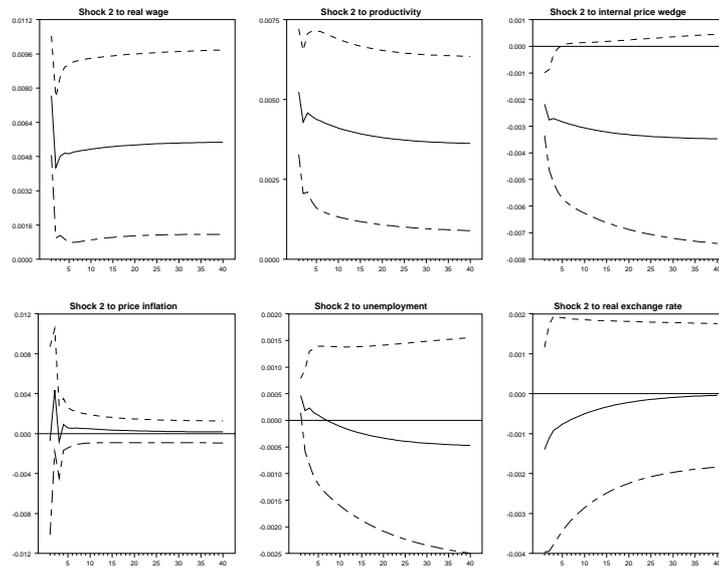


Figure 6: Impulse responses after a shock to productivity

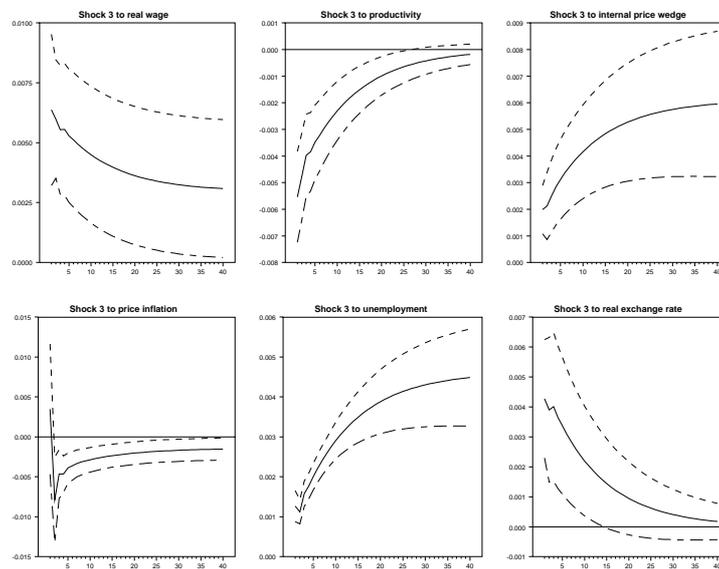


Figure 7: Impulse responses after a shock to nominal wages

In addition to the trend of foreign origin and the trend emanating from shocks to domestic productivity, we have two more shocks to interpret. The impulse responses corresponding to the third trend, interpreted as arising in the labour market, is presented in figure 7. In the graphs, we see that both unemployment and real wages increase in response to this shock. A rise in nominal wages captures these two effects, since an increase in nominal wages leads to higher real wages and thus higher costs for firms, which may cause rising unemployment. A nominal wage shock may arise from the wage setting behaviour in the economy if strong labour unions can achieve large increases in nominal wages. A possible explanation of the negative reaction of productivity in the short run is that if decisions and investments are based on predictions concerning the firm's costs, including wage costs, a sudden rise in wages can make previous investment unproductive and lead to a fall in productivity. Note that after the initial fall, productivity rises back to the original level in close connection with the rising unemployment. The reaction of the producer price inflation is harder to explain. A rise in nominal wages should lead to a rise in inflation, but the reaction of producer price inflation is the opposite. The reaction of the internal price wedge indicates that consumer prices are increasing after this shock. This also explains the temporary effect on the real exchange rate, where the Swedish price level causes the real exchange rate to appreciate.

Finally, the impulse response functions after a shock to the fourth trend are presented in figure 8. We consider this to be a shock to  $p^{c/y}$ , the internal price wedge. The identifying restrictions imply that a shock to this trend does not have a long-run impact on import prices, productivity or unemployment. The interpretation is supported by the empirical estimates of the impulse response functions, where the effect on the price wedge is significant. The negative response of this variable indicates that producer prices have increased, which can be the result of a higher degree of monopoly in the manufacturing sector. Increased producer prices result in a lower real product wage. A shock to this trend causes a short-run effect on unemployment. This is in line with economic theory since a higher degree of monopoly in the economy should reduce production and unemployment. Producer price inflation initially increases but quickly moves back to its initial level.<sup>31</sup>

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<sup>31</sup>When we impose an identifying zero-restriction on unemployment in the  $\Upsilon$ -matrix in section 6.1, an

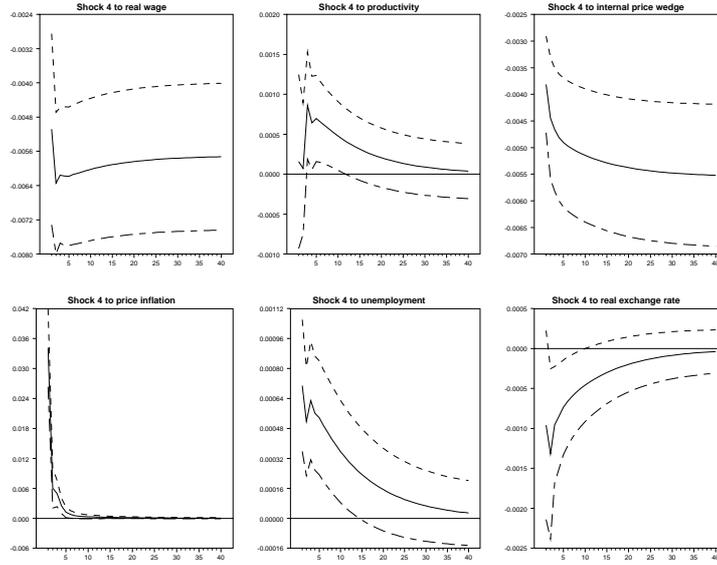


Figure 8: Impulse responses after a shock to the internal price wedge

As seen in the impulse response analysis, real wages react to all of the four common trends. A positive shock to productivity increases real wages, as expected from economic theory, and a rise in nominal wages leads to a rise in the real wage as well. Following an increase in prices, occurring from either import prices or producer prices, the real wage falls, which is natural. Unemployment, on the other hand, does not react strongly to productivity shocks. Historically, we have not seen any tendency for unemployment to rise or fall in connection with productivity growth, and the result in the impulse response analysis is in line with this observation. Nor do shocks to import prices influence unemployment. The two shocks that affect unemployment are the labour market shock and the shock to the degree of monopoly in the producer sector. A rise in nominal wages has a strong effect on unemployment and leads to a higher unemployment level. If the degree of monopoly in manufacturing increases, this also induces higher unemployment. We interpret this as further signs of rigidities present in the Swedish labour market, since, without rigidities, the shocks to the labour market and the monopoly situation otherwise should have resulted in changes in wages instead of in unemployment.

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additional zero-restriction is imposed on producer price inflation for the fourth shock. This additional restriction is a result of algebraic transformations made when calculating the identification matrix  $B$ .

## 7 Conclusions

We study the wage formation in Sweden between 1982 and 2002 and the relation between wages and unemployment in an empirical investigation using the cointegrated VAR approach. We estimate the empirical wage determination process based on a theoretical model with wage bargaining between trade unions and firms, where the real wage depends on productivity, unemployment and the price wedge between consumer and producer prices. In addition to the wage determination equation, we estimate a dynamic price equation, where price inflation depends on unemployment and the real exchange rate. Compared to the results in Jacobson et al. (1997), who also examine the Swedish labour market, we find a relation between the real wage and unemployment, which is not found in Jacobson et al. (1997). In a recent study, Friberg (2004) finds a negative relation between wages and unemployment, but he does not find an explicit link between productivity and wages, contrary to what is expected from economic theory. Here, we find a richer pattern of interaction where real wages depend on both productivity and unemployment.

The negative relation between the real wage level and the unemployment level implies that higher unemployment leads to less pressure on the wage level. The elasticity between the variables is estimated to be 0.76, which can be compared to Pétursson and Sløk (2001), who estimate the elasticity to be 0.3 in the Danish labour market. The elasticity we find for Sweden is also higher than Bårdsen et al. (2001) find for Australia and Bårdsen and Fisher (1999) find for the United Kingdom. It thus seems as if Swedish wages react more strongly to changes in unemployment than wages in Denmark, Australia and the United Kingdom.

We also find evidence of rigidities in the Swedish labour market. With no rigidities in the market, productivity and unemployment would not be influenced by nominal variables or by nominal shocks in the economy, but our results indicate that they are affected by such shocks. The estimate of the price equation shows that price inflation and unemployment are negatively related as in the traditional Phillips curve. We find that the relation between unemployment and prices is stronger than the relation between wages and prices. This implies that fluctuations that would normally appear in the wage level appear in the

unemployment level instead. In the common trends analysis, unemployment reacts strongly to shocks to nominal wages and to shocks to the internal price wedge, which is also a sign of rigidities. In addition, both productivity and unemployment adjust to equilibrium deviations in the real wage relation, an adjustment that would not be present without rigidities.

The analysis suggests a large influence of trade unions in the wage formation process. As real wages are found to closely follow the development in consumer prices, and not the development in producer prices, it seems that trade unions have the greater influence in the wage bargaining process. The estimated results also imply that there is a large deviation between the development in real wages and the development in productivity. The one-to-one relationship implied by economic theory is not found in the data. Instead, the estimates indicate that a rise in productivity leads to a real wage rise more than twice as high. This is at variance with previous studies concerning Sweden, where a one-to-one relation between real wages and productivity generally emerges (Jacobson et al. (1998), Jacobson et al. (1997)). However, in these studies, the analysis covers other time periods. A relation that diverges from the homogenous one is found for Australia in Bårdsen et al. (2001) and for the Euro-area in Juselius (1999).

The overall picture is that Swedish real wages have grown too fast compared to productivity throughout the period studied. A high wage level implies large costs for firms and is likely to be one of the causes behind the high unemployment level. Due to rigidities in the labour market, the high unemployment is persistent. This study thus supports the hypothesis that institutional factors in the Swedish labour market give rise to higher unemployment. Hence, reforms that contribute to reducing these rigidities are likely to be a fruitful way of lowering unemployment.

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