Permanent Shocks and Spillovers:

A Sectoral Approach Using a Structural VAR*

By

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Abstract

In an empirical two-sector model, it is investigated if permanent shocks originating from one sector have an effect on production in the other sector. The VAR contains three variables: employment in the private sector, and value added in industry and in the rest of the private sector. I found that shocks originating from one sector have no effect on production in the other sector, neither in the short, nor in the long run. The only common source of fluctuations in sectoral production is shocks to the labor trend.

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1. Introduction

Small open economies can be divided into one sector with a high degree of international dependence, and another sector with production directed towards the home market. In Sweden, the export share of production in the industrial sector is more than fifty percent, while in the rest of the private sector it is less than ten percent. In the latter, the value added is twice as large as in industry, reflecting the importance of the home market sector. Typically, productivity growth in the industrial sector is higher than in the home market sector, a difference that affects the development of relative prices of the sectors.

How is production in one sector related to production in the other? Do shocks originating from one sector have any effect on production in the other sector, and if such effects are present, how important are they? These questions are relevant both in business cycle analysis, and when studying growth perspectives in the two sectors. One would expect that there exist spillover effects, since supply shocks will alter the relative prices of the goods produced in each sector. Also, it is usually believed that fluctuations in the exporting sector influence the business cycle dynamics of the home-market production.

I use a structural VAR to answer the question about the relation between production in the two different sectors. This is a data-oriented approach, where the general dynamic pattern in data is reflected in the empirical model. A few identifying assumptions have to be kept as maintained hypotheses, and other overidentifying restrictions are testable. A model, similar to the one used in Blanchard and Quah (1989) is extended to two production sectors. Implications from a theoretical model are used as identifying assumptions on the long run behavior.

The economic model used for identification is presented in Section 2. Section 3 gives details about the econometric method and about specification. Section 4 describes the data and presents the results from the estimation of the unrestricted VAR. Section 5 contains the estimation results when structure is imposed. Some concluding remarks are made in Section 6.

2. The Model

In this section a simple two-sector model of output and employment determination is outlined. The model draws on Fischer (1977) and Blanchard and Quah (1989), viewing fluctuations in a standard Keynesian framework and provides an economic interpretation of both short-run and long-run responses to real and nominal disturbances. Implications from the model are used as identifying assumptions on the long run responses and for the choice of variables in the empirical analysis.

The theoretical model is the following:

$$y_t^i = m_t - p_t^i, \qquad \textit{i=1,2} \qquad (1)$$

$$y_t^i = g^i n_t^i + q_t^i \qquad \textit{i=1,2} \qquad (2)$$

$$p_t^i = w_t^i - q_t^i + (1 - g^i) n_t^i, \qquad \textit{i=1,2} \qquad (3)$$

$$w_t = w_t | (E_{t-1} n_t = l_{t-1}). \qquad (4)$$

The variables y_t^i, p_t^i, w_t^i, n_t^i, and q_t^i are the logs of output, prices, nominal wages, employment and the productivity level in each sector, respectively. The variables m_t, w_t, n_t and l_t denote the log of money supply, the aggregate nominal wage, employment in both sectors, and labor force, respectively. Finally, g^i is a parameter reflecting the returns to scale in each sector. Constants are normalized to zero.

Equation *(1)* describes demand for good *i*, consistent with a Cobb-Douglas utility function. Equation *(2)* is the production function that relates output in a sector to employment and sector-specific productivity, allowing for non-constant returns to scale, reflected in the parameter g^i[1]. Equation *(3)* is the zero-profit condition[2] that relates the output prices to wages, productivity and employment. Equation *(4)* describes wage setting in the economy. Wages are determined one period ahead, and are set so that expected total employment will

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[1]Since there is only one production factor, labor, in the theoretical model, the possibility of decreasing returns to scale incorporates a CRS technology with two production factors, where the other factor, capital, is fixed.

[2]Equation *(3)* is obtained by substituting the expression for y_t^i in *(2)* into the zero-profit condition $p_t^i + y_t^i = w + n_t^i$.

equal labor force. This wage setting behavior is consistent with a trade union that cares about both employed and unemployed members and therefore, outsiders influence the wage setting to some extent. The sectoral nominal wages are equal since all individuals in the labor force could be employed in both sectors[3].

Closing the model, and characterizing each sector's response to shocks, I need to specify how the exogenous variables, productivity $\left(q_t^i\right)$, labor force $\left(l_t\right)$, and nominal money $\left(m_t\right)$, evolve over time. I assume that they follow random walk with drifts:

$$l_t = m^l + l_{t-1} + e_t^l \qquad (5)$$

$$q_t^i = m^i + q_{t-1}^i + e_t^i \qquad i=1,2 \qquad (6)$$

$$m_t = m^d + m_{t-1} + e_t^d, \qquad (7)$$

where e_t^l, e_t^i and e_t^d denote shocks to labor force, sector specific productivity and nominal demand respectively, which all are assumed to be mutually independent white noise.

Substituting *(5)-(7)* into *(1)-(4)* yields the following solution for employment and sectoral production, n_t and y_t^i:

$$n_t = e_t^{(d)} + m^{(l)}t + \sum_{j=1}^{t} e_{t-j}^{(l)} \qquad (8)$$

$$y_t^{(i)} = g^{(i)}e_t^{(d)} + g^{(i)}(m^{(l)}(t-1) + \sum_{j=1}^{t} e_{t-j}^{(l)}) + m^{(i)}t + \sum_{j=0}^{t} e_{t-j}^{(i)} \qquad i=1,2 \qquad (9)$$

Nominal money $\left(e_t^d\right)$ has only short-run effects on production and employment. In the long run, employment follows the labor trend, see equation *(8)*, and production follows both the sector specific productivity trend and the labor trend, equation *(9)*. There is no spillover between the sectors since sectoral production is affected by sector-specific productivity shocks $(e_t^i$, where $i \neq j)$ only. This implication arises from the specific parameterizations of the demand and production functions in equations *(1)* and *(2)*. If, for example, substitution

[3]This assumption is not crucial and sectoral wages could be allowed to differ, for example by assuming exogeneous labor supply to each sector. In the empirical analysis it is not possible to separate effects from two different labor trends, since data on sector-specific labor supply is not available.

between goods is allowed for, or, as in Blanchard and Quah (1989), a productivity term is introduced ad hoc in the demand function, this implication will not hold.

The model shows that it is not enough to include only productivity shocks in the two sectors, and that shocks to a stochastic labor trend affect both sectors. Therefore I also need to include employment in the empirical model for sectoral production, to distinguish between the sector specific productivity shocks and the common labor force shock.

With three variables, employment and production in each sector, three shocks of the four in the model could be identified. The temporary demand shock, e^d, will not be analyzed here. Instead I will make a loose interpretation of the three supply shocks, (e^1, e^2, e^l), in the theoretical model, as being a mixture of supply and demand shocks, having permanent effects. In applications where permanent shocks are labeled supply shocks and temporary shocks demand shocks it is often found that permanent shocks are important also in the short run. On the other hand, demand shocks defined in another way could have persistent effects, being important in the long run, see for example Assarsson and Jansson (1996), where cyclical shocks are found to have permanent effects on unemployment. In the empirical analysis I will study the effects for different time horizons from permanent shocks of which two are sector specific (i.e. the productivity shocks) and one is common for both sectors (i.e. the labor shock).

3. Specification, Identification and Estimation

For the purpose of analyzing the effects from sector specific shocks on production, I propose a structural VAR model, consistent with the theoretical considerations above. In equations *(8)-(9)*, production and employment are assumed to be I(1) variables, but not cointegrated, in which case a VAR in differences would be appropriate:

$$B(L)\Delta x_t = d + h_t, \qquad (10)$$

where x_t is a *3 ´1* vector containing the endogenous variables, (the log of) employment and production in Sector 1 and two, respectively, $x_t = \begin{bmatrix} n_t & y_t^1 & y_t^2 \end{bmatrix}'$. $B(L)$ is a matrix polynomial in the lag operator of order p, $B(L) = I_n - \sum_{j=1}^{p} B_j L^j$, where $L^j z_t = z_{t-j}$, and it is assumed to have all its roots outside the unit circle, d is a 3×1 vector containing constants, and h_t is a vector of zero-mean, i.i.d. disturbances and $E(h_t h_t') = \Sigma$, a positive definite matrix.

The structural disturbances, i.e. the exogenous shocks in equations *(5)* to *(7)*, e_t, are assumed to be linearly related to the reduced form (forecasting) errors in *(11)*, $h_t = Ce_t$. The structural innovations e_t are also assumed to be i.i.d., and their variances are normalized to unity, to facilitate comparisons of the effects from the shocks, $E(e_t e_t') = I$. Hence, we have $\Sigma = E(h_t h_t') = E(Ce_t e_t' C') = CE(e_t e_t')C' = CC'$.

The Vector Moving Average, (VMA), or Wold representation, relates the endogenous and the exogenous variables to each other:

$$\Delta x_t = B(1)^{-1} d + B(L)^{-1} h_t = D(1)m + D(L)e_t, \qquad (11)$$

where $m = C^{-1} d$ and

$$D(L) = B(L)^{-1} C = \sum_{j=0}^{\infty} D_j L^j \ . \qquad (12)$$

The exogenous shocks are related to the levels of the variables according to the theoretical model *(8)*-*(9)*, which can be seen by reformulating the model. The *D(L)* polynomial in *(11)* can be decomposed as $D(L) = D(1) + D^*(L)\Delta$, where $D_j^* = -\sum_{i=j+1}^{\infty} D_i$, see Engle and Granger (1987). The levels are obtained by recursive substitution in the VMA model $\Delta x_t = D(1)(m + e_t) + D^*(L)\Delta e_t$, where the initial values are set to zero, resulting in:

$$x_t = D(1)\left(mt + \sum_{i=1}^{i=t} e_{t-i} \right) + D^*(L)e_t.$$

(13)

Since the expression within the parenthesis is a random walk vector, we can equivalently express *(13)* in terms of the trends:

$$x_t = D(1)t_t + D^*(L)e_t, \tag{14}$$

where $t_t = m + t_{t-1} + e_t$. From *(14)* we see that the long run effects from the permanent shocks to the stochastic trend in $t_t = [l_t \quad q_t^1 \quad q_t^2]$ are measured by the coefficients in the *D(1)* matrix, and that *D*(L)* measures the short-run effects from the same shocks.

The implications from the economic model can now be used as identifying assumptions on the long run responses to the shocks, which are measured by the coefficients in the *D(1)* matrix. The economic model, *(1)-(4),* implies that, in the long run, the common labor force trend and the sector-specific productivity trends drive production in each sector, and that employment is driven by the labor force trend only. The long run impact matrix, *D(1)* will, according to the model, have the following form:

$$\begin{bmatrix} n_t \\ y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} d(1)_{11} & 0 & 0 \\ d(1)_{21} & d(1)_{22} & 0 \\ d(1)_{31} & 0 & d(1)_{33} \end{bmatrix} \times \begin{bmatrix} l_t \\ q_t^1 \\ q_t^2 \end{bmatrix} + D^*(L)e_t.$$

Three identifying restrictions have to be introduced on *D(1),* in order to achieve exact identification of the shocks. It turns out that the shocks could be identified in two different ways. Identification can be achieved by imposing the following (three zero) restrictions on the long run responses:

$$\begin{bmatrix} n_t \\ y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} d(1)_{11} & 0 & 0 \\ d(1)_{21} & d(1)_{22} & 0 \\ d(1)_{31} & d(1)_{32} & d(1)_{33} \end{bmatrix} \times \begin{bmatrix} l_t \\ q_t^1 \\ q_t^2 \end{bmatrix} + D^*(L)e_t. \tag{15}$$

I am then able to test if there is long run spillover from Sector 1 to Sector 2, i.e. if $d(1)_{32}=0$, given that the sector specific trend in Sector 2 has no long run effect on production in Sector 1.

Alternatively, another set of zeros can be imposed on the long run response, so that

$$\begin{bmatrix} n_t \\ y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} d(1)_{11} & 0 & 0 \\ d(1)_{21} & d(1)_{22} & d(1)_{23} \\ d(1)_{31} & 0 & d(1)_{33} \end{bmatrix} \times \begin{bmatrix} l_t \\ q_t^1 \\ q_t^2 \end{bmatrix} + D^*(L)e_t \,. \qquad (16)$$

Here, I can test long run spillover from Sector 2 to Sector 1, given that Sector 1 shocks have no long run effect on production in Sector 2, i.e. if $d(1)_{23}=0$.

Technically, I need to specify the matrix C that connects the economically interpretable structural disturbances e_t with the forecasting errors in the VAR-model, h_t. The C-matrix should both diagonalize Σ and, at the same time, satisfy the identifying restrictions. Once C is known, the parameters in the *D(L)*-polynomial can be recovered from estimation of the reduced reduced form polynomial *B(L),* using *(12).* The computation of the C-matrix is simple in this case, since the long run matrix, *D(1),* in line with the interpretation of the economic model, can be written as a lower triangular matrix.

D(1) is calculated from a Cholesky-decomposition of $D(1)D(1)' = B(1)^{-1}CC'B(1)^{-1'} =$ $= B(1)^{-1}\Sigma B(1)^{-1'}$, where $B(1)^{-1}$ and Σ are known from the estimated VAR-model. Again using *(12)* the C-matrix can be computed from $C = B(1)D(1)$. The C^{-1}-matrix has the property that it diagonalizes Σ, since $E(e_t e_t') = C^{-1}\Sigma C^{-1'} = C^{-1}CC'C^{-1'} = I$.

4. Data and the unrestricted VAR

The variables, employment and production in the two sectors, in the VAR-model *(10)* are assumed to be I(1) and not cointegrated. Results from tests of integration and cointegration are reported at the end of Section 4.2.

The data set consists of semi-annual seasonally non-adjusted observations, 1970-1995, on employment measured in hours in the private sector, value added in industry, (ISIC 2 and 3), and in the rest of the private sector, other business, (ISIC 1, 4-9). Sector 1 is essentially the manufacturing industry, and Sector 2 is the rest of the private sector. I use National Accounts data as stored in the KOSMOS[4]-database, the sectors corresponding to the two private production sectors in KOSMOS. All variables are in logarithms.

4.1 Preliminary data characteristics

The logs of employment, *n*, and production in the two sectors, y^1 and y^2 are plotted in Figure 1. All variables are trending, while the cyclical swings are more pronounced for employment and production in Sector 1. There is seasonal variation in the data that needs modeling. An informal way of judging whether the seasonal variation is better described by a deterministic or a stochastic model is to plot the data as annual observations, (one time series for each semi-annual observation), see Franses (1990). If the two curves cut each other several times, it can indicate that the seasonal variation is better described by a stochastic model. If they do not cut each other, a model with deterministic dummies is worthy testing. As can be seen from Figure 2, deterministic dummies could be reasonable for *n*, while a stochastic model seems reasonable for y^2. I am not able to rely on any formal test for unit roots at the seasonal frequency[5], and there is a risk of overdifferentiation if I use a stochastic model, and the low number of observations makes a periodic model unsuitable. I therefore estimate the VAR with deterministic dummies, refraining from modeling changing seasonals.

4.2 Estimation of the unrestricted VAR

The number of lags in the VAR model is determined from information criteria and mis-specification tests. The results are reported in Table 1. The Schwartz criterion picks out two and the Hannan-Quinn, (HQ), criterion three lags. I choose three lags since it is the smallest

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[4]This study should also be used as an input in another project that aims at comparing the results in this paper with simulation experiments using KOSMOS, a large scale econometric model at the National Institute of Economic Research, Sweden.

[5]I calculate the HEGY test, taking into account that semi-annual data do not have imaginary roots. Unfortunately, there are no critical values for semi-annual data in Hylleberg et. al. (1990). If I use their critical values for root -1, the tests indicate that the null of seasonal unit roots could be rejected for all variables at the 10 per cent level. The correct critical values are probably lower than in Hylleberg et. al. The critical value in the Dickey Fuller distribution tends to increase with the number of nuisance parameters, and since the auxiliary regression with semi-annual data has a lower number of such parameters, I am probably able to reject the null of seasonal unit roots at a lower *p*-level than ten per cent.

number of lags where autocorrelation in the residuals disappears, both according to the LM and to the Portmanteau test.

Table 1. Information criterion and *p*-values for multivariate mis-specification tests

Nr. of lags	1	2	3	4	5	6	7	8
Schwarz	23.91	24.48*	24.40	23.92	23.06	23.44	23.27	22.73
HQ	24.29	25.09	25.30*	24.99	24.96	24.97	25.03	24.72
LM(2)	0.000	0.054	0.167	0.228	0.073	0.231	0.675	0.836
LM(4)	0.000	0.010	0.110	0.249	0.384	0.419	0.585	0.668
Portm(15)	0.000	0.498	0.709	0.680	0.045	0.070	0.010	0.001
Normality	0.500	0.137	0.412	0.170	0.143	0.237	0.098	0.031

Note: *LM(2)*, *LM(4)*, and *Portm(15)* are tests for error autocorrellation. All test statistics are calculated with PcFiml 8.0. * denotes the preferred number of lags for each information criterion.

In Table 2, likelihood ratio tests are reported. The first test is a sequential test for the number of lags, and the second is a direct test against eight lags. According to both LR tests, three lags seem to be reasonable.

Table 2. LR-test for lag length

H_0	1 in 2	2 in 3	3 in 4	4 in 5	5 in 6	6 in 7	7 in 8
p-value	0.000	0.004	0.668	0.138	0.162	0.194	0.903
H_0	1 in 8	2 in 8	3 in 8	4 in 8	5 in 8	6 in 8	7 in 8
p-value	0.000	0.088	0.337	0.265	0.404	0.603	0.903

Note: “1 in 2” means that the null is 1 lag and the alternative is 2 lags.

Chow-tests, (not reported here) indicate a change in the parameters in 1993. Introducing the log of the nominal effective exchange rate into the model, as an exogenous variable, the parameters stop varying over time. None of the estimated model parameters are affected by the inclusion of the exchange rate, regardless of whether it is allowed to have long run effects or not. The reaction of the Chow test seems to be due to a temporary turbulence at the change of exchange rate regimes. Here, only results from an estimated model excluding the exchange rate are reported.

The modelling approach presumes that the variables are I(1) and not cointegrated. Alternatively there could be two cointegrating vectors among three variables, excluding the unrealistic possibility that all variables would be stationary, (cf. Figure 1). However, there is only one economically meaningful long run relation that can appear in this dataset, namely a cointegrating relation between production in the two sectors. As can be seen from Figure 3, relative production does not look like a stationary variable. Other combinations of the variables, total employment and production in each sector or in both sectors, are not interpretable, but they may be loosely related to labor productivity. Since labor productivity is trending, other combinations of the variables than sectoral production, are not candidates for a cointegrating relation.

The number of cointegrating relations is tested for, using the max eigenvalue and the trace tests. The results reported in Table 3 indicate two cointegrating relations, i.e. one common trend. This result is surprising. As could be seen from Figure 4, where the deviation from the first cointegrating relation is plotted, it does not look like a stationary variable. This raises the question about the small sample properties of the tests. Here, no attempt is made to investigate this issue further, but we can note that Jacobsson, Vredin and Warne (1997) obtain the result that the empirical distribution of the trace statistic is skewed to the right of the asymptotic distribution.

Table 3. Cointegration tests

H_0	*Max(T)*	*Max(T-nk)*	*Trace(T)*	*Trace(T-nk)*
$r = 0$	24.42**	19.93*	43.35***	35.58**
$r \leq 1$	16.47**	13.44*	18.93**	15.45**
$r \leq 2$	2.46	2.01	2.46	2.01

Note: A * denotes significant at the 10 percent level, ** at the 5 percent level, and *** at the 1 percent level, respectively.

To check if production in the sectors are related to each other in the long run, I test, conditional on one cointegrating vector, if employment could be excluded. The *p*-value for this test is zero, and therefore we can conclude that a formal test does reject the hypothesis that the production levels are tied to each other in the long run.[6]

To check if sectoral productions is related in the long run when a trend is allowed for in the cointegrating relation, I re-estimate the model including a deterministic trend. Here too, the *p*-value is zero for the hypothesis of exclusion of employment, so sectoral production is not stationary around a deterministic trend, according to a formal test.

From the discussion above about the small sample properties of the tests for the number of trends, together with the result that sectoral production probably is non-stationary, there seems to be no strong empirical evidence against the assumption of three separate stochastic trends, as in the theoretical model.

5 Estimation results

Two alternative models, with different identifying assumptions, are estimated. In the first model, the productivity trend in Sector 1 is identified by the assumption that there is no spillover from Sector 2 to Sector 1, see equation *(16)*. In the alternative model, the identifying assumptions are alternated, so that there is no spillover from Sector 1 to Sector 2, see equation *(17)*. In the short run, all shocks are allowed to affect the variables.

The standard errors are computed from 5000 parametric bootstrap replications. The starting values are generated from the estimated distribution, and new errors are drawn from a normal distribution, with a covariance matrix equal to the one estimated.

[6]The *p*-value is also zero for the test of stationarity of relative production, i.e. employment excluded and equal coefficients with different signs on production in the cointegrating relation.

5.1 The long run effects

For the model *(16)* the estimated long run effects are the following:

$$\begin{bmatrix} n_t \\ y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} \underset{(0.00344)}{0.0188^{***}} & 0 & 0 \\ \underset{(0.00879)}{0.0174^{**}} & \underset{(0.00519)}{0.0313^{***}} & 0 \\ \underset{(0.00326)}{0.0055^{*}} & \underset{(0.00280)}{0.00323} & \underset{(0.00218)}{0.0104^{**}} \end{bmatrix} \times \begin{bmatrix} \underset{(0.179)}{-0.177} + l_{t-1} + e_t^l \\ \underset{(0.211)}{0.318} + q_{t-1}^1 + e_t^1 \\ \underset{(0.463)}{0.903} + q_{t-1}^2 + e_t^2 \end{bmatrix} + D^*(L)e_t \, .$$

Where * denotes significantly different from zero at the ten per cent level, ** at the five per cent level, and *** at the one per cent level. Standard errors are shown in parentheses.

The common labor trend has significant effects on all variables. Sector 1 shocks have significant impact on production in Sector 1, but not on production in Sector 2. Sector 2 shocks significantly affect production in that same sector.

The estimation results for model *(16)* are:

$$\begin{bmatrix} n_t \\ y_t^1 \\ y_t^2 \end{bmatrix} = \begin{bmatrix} \underset{(0.00355)}{0.0188^{***}} & 0 & 0 \\ \underset{(0.00886)}{0.0174^{**}} & \underset{(0.00474)}{0.0299^{***}} & \underset{0.00860}{0.00925} \\ \underset{(0.00322)}{0.0055^{*}} & 0 & \underset{(0.00243)}{0.0109^{***}} \end{bmatrix} \times \begin{bmatrix} \underset{(0.177)}{-0.177} + l_{t-1} + e_t^l \\ \underset{(0.501)}{0.037} + q_{t-1}^1 + e_t^1 \\ \underset{(0.418)}{0.956} + q_{t-1}^2 + e_t^2 \end{bmatrix} + D^*(L)e_t \, .$$

The significant coefficients estimates are almost identical to those in the first model.

Thus, irrespective of the identifying assumptions, the results indicate that there is no spillover between the sectors. We can conclude that growth in one sector is independent of growth in the other sector in the long run, in the sense that only the own sector shocks matter. Furthermore, all other coefficients are significantly different from zero, and they have the expected signs.

5.2 The estimated trends

An alternative way of evaluating the relevance of the interpretation of the trends is to plot them together with the theoretical variables they are supposed to represent. This is done in Figure 5. In the comparisons the estimated labor trend is multiplied by its long run effect on employment, and the two productivity trends are multiplied by their long run coefficients on production. This means that the trends are normalized in such a way that a one per cent increase in the labor and productivity trends changes employment and production by one per cent in the long run. The levels of the trends are adjusted so as to coincide with the variables they are plotted together with in 1971.

Private employment is used in the estimation and labor force data are only available for the whole economy. Since 1970 private employment's share of total employment has decreased steadily, and the growth in total employment in the economy is attributed to growth in government employment. In the comparisons, labor force data are adjusted according to the downward trend in private employment, assuming that the share of the ”effective” labor supply to the private sector follows the private sector's share of total employment. Furthermore, the number of persons in the labor force is converted into hours by using the average hours of work per employed. As can be seen from the graph, the estimated labor trend captures the general behavior of this adjusted measure of labor force. In Figures 5.b and 5.c the estimated productivity trends are plotted together with labor productivity in the two sectors. Here too, the estimated trends capture some relevant aspects of the behavior of the observed labor productivity.

5.3 Short run effects

The $D^*(L)$ polynomial, in *(15)*, represents the short run effects from the shocks. Figure 6 shows the impulse responses[7] for the first 20 years, and the 95 per cent confidence bands.

[7]The short run responses are almost identical in the two alternative models. The rest of the estimation results presented, are based on the identifying assumptions in the first model. The overidentifying restriction of no spillover is, for computational reasons, not imposed.

Figure 6 shows that a *labor shock* results in a permanent increase in employment and production. The adjustment is gradual and peaks are reached at 2-3 years, and it takes six years to reach the new long run levels. Unemployment is assumed to be unaffected by shocks to the labor trend in the long run, since employment follows the labor trend in the long run. Hence, the gradual response in employment implies a temporary increase in unemployment. The short run response in production in Sector 1 is greater than the long-run response, while the short-run response in Sector 2 is smaller than the long-run response, although not significantly. This difference in the point estimates could be explained by the two sectors having different marginal productivity, reflecting differences in the capital-labor ratio, changes in relative employment, and different wage setting behavior.

As expected, a S*ector 1 shock* significantly increases production in Sector 1, c.f. Figure 7. The adjustment is gradual and it takes three years to reach the new long run level. Employment decreases initially, which implies a temporarily increased unemployment. The response in production in Sector 2 is not significantly different from zero. We can conclude that Sector 1 shocks do not affect production in Sector 2 at any time horizon.

A S*ector 2 shock* increases production in Sector 2, as expected, cf. Figure 8. Generally, the responses to a Sector 2 shock are not smooth, reflecting that not all of the seasonal variation is accounted for in the estimations, cf. the discussion in section 4.1. The short-run effect on production in Sector 1 is only significantly different from zero in the first period, and we can conclude that Sector 2 shocks do not affect production in Sector 1 after the first period. Compared to the case of a Sector 1 shock, there is a difference in the response of employment, which here increases temporarily, suggesting a lower unemployment, although the question of significance is blurred by the aforementioned misspecification of the seasonal component.

6 Concluding remarks

In an empirical two-sector model, I test if permanent shocks originating from one sector have any effect on production in the other sector. In general, one expects that sector specific supply shocks have spillover effect on production in the other sector, through changes in the relative output price. I found no empirical evidence of spillover effects from sector specific shocks, neither in the short nor in the long run. The only common source of fluctuation in sectoral production is shocks to the labor force trend, which affect production in both sectors. The results also hint that there may be a difference in the short run adjustment of employment to sector specific shocks. As an example, the temporary decrease in employment from a Sector 1 shock, could be explained by nominal rigidities, where demand is ”too low” to match the increased production, and instead it leads to a temporary increase in unemployment. The same factors, as for Sector 1 shocks, could not explain the estimated response in employment to a Sector 2 shock, since employment in this case is increased. One reason could be differences in the wage and price-setting behavior, variables that are not included in the estimations. Also, the results indicate short run differences in the response of sectoral production to labor shocks.

It turns out that the introduction of a variable capturing shocks to both sectors is important. If one asks the same question using a system containing only sectoral production and no employment, one will find spurious evidence of spillovers in both directions (since production in the two sectors would then Granger-cause each other).

Only shocks with permanent effects are included in the analysis, and it may be interesting to return to the issue about spillovers being able to separate sector specific permanent (supply) and transitory (demand) shocks to see if the results in this paper then hold.

References

Assarsson, B., and Jansson, P. (1996), “Unemployment persistence: the Case of Sweden”, forthcoming in *Applied Economics Letters*

Blanchard, O.J., and Quah, D. (1989), ”The Dynamic Effects of Aggregate Demand and Supply Disturbances”, *The American Economic Review*, vol. 79, no. 4, 654-673

Dolado, J.J., and Lopez-Salido, J.D. (1996), ”Hysteresis and Economic Fluctuations (Spain, 1970-94)”, Discussion Paper No. 1334, Center for Economic Policy Research

Engle, R.F., and Granger, C.W.J. (1987),”Co-Integration and Error Correction: Representation, Estimation and Testing”, *Econometrica*, 55:251-276

Fischer, S. (1977), ”Long-Term Contracts, Rational Expectations, and the Optimal Money Supply Rule”, *Journal of Political Economy*, 85, no.1, 191-205

Franses, P.H. (1990), “A multivariate approach to modeling univariate seasonal time series”, Econometric Institute report no. 9101, Erasmus University, Rotterdam

Hylleberg, S., Engle, R.F., Granger, C.W.J., and Yoo, B.S. (1990), ”Seasonal Integration and Cointegration”, *Journal of Econometrics*, 44, 215-38

Jacobsson, T., Vredin, A., and Warne, A., (1997), ”Common Trends and Hysteresis in Scandinavian Unemployment”, *European Economic Review*, 41, 1781-1816

Fig. 1a Employment in the private sector

Fig. 2a Employment in the private sector

Yearly observations

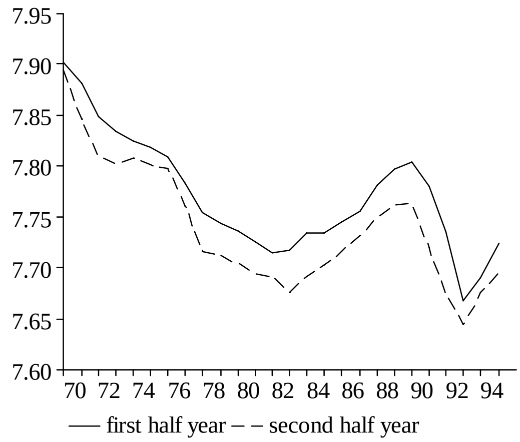


Fig. 1b Production in Sector 1

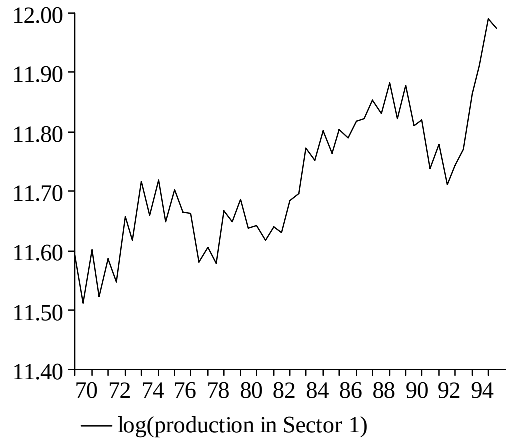


Fig. 2b Production in Sector 1

Yearly observations

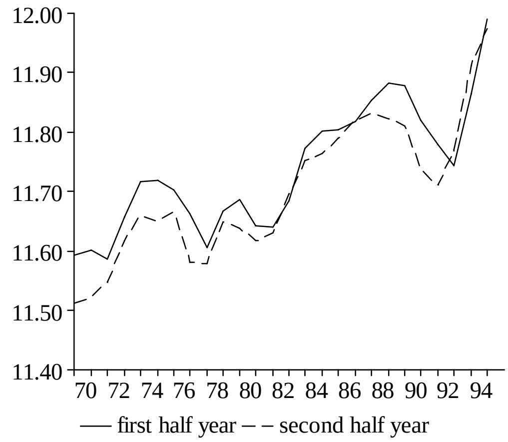


Fig. 1c Production in Sector 2

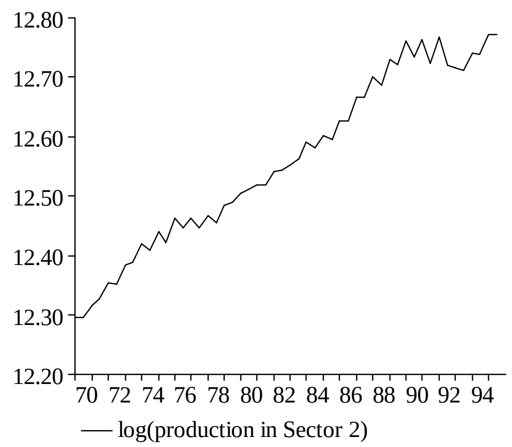


Fig. 2c Production in Sector 2

Yearly observations

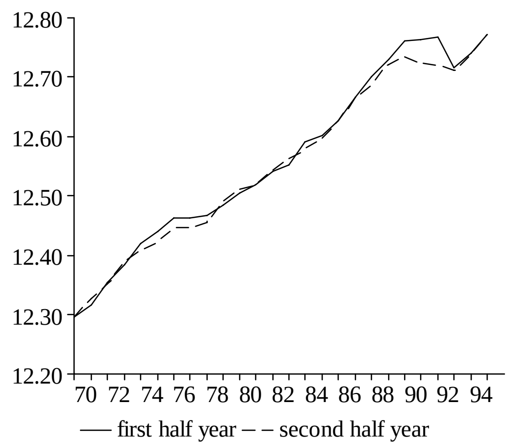


Fig. 3 Ratio of production in Sector 1 and Sector 2

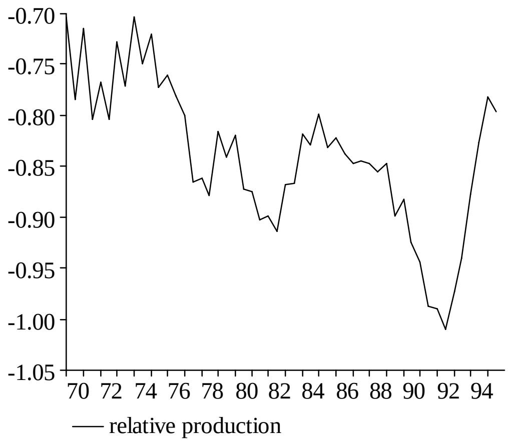


Fig. 5a Adjusted labor force and the estimated labor trend

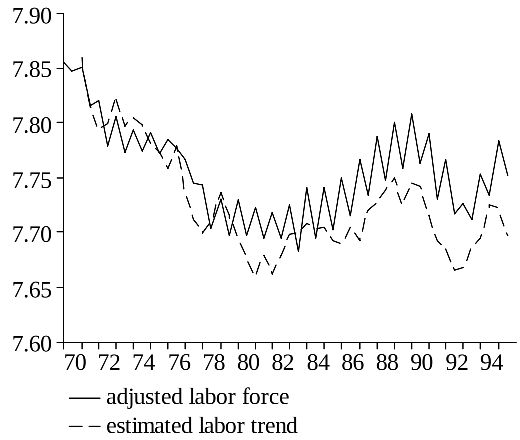


Fig. 4 Deviations from the first cointegrating relation

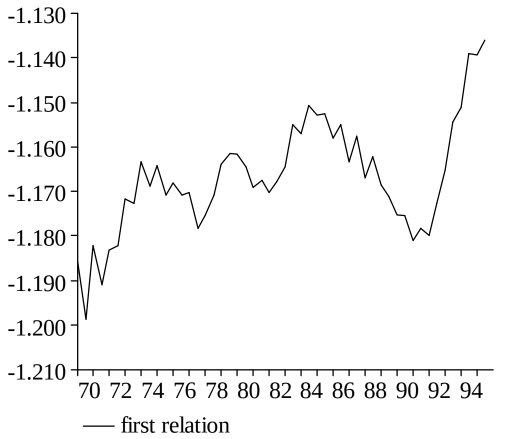


Fig. 5b Labor productivity in Sector 1 and the estimated productivity trend

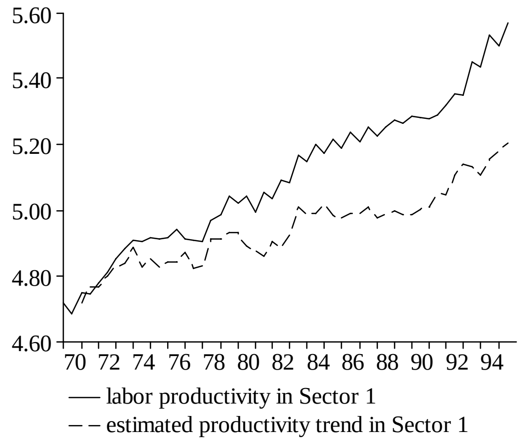


Fig. 5c Labor productivity in Sector 2 and the estimated productivity trend

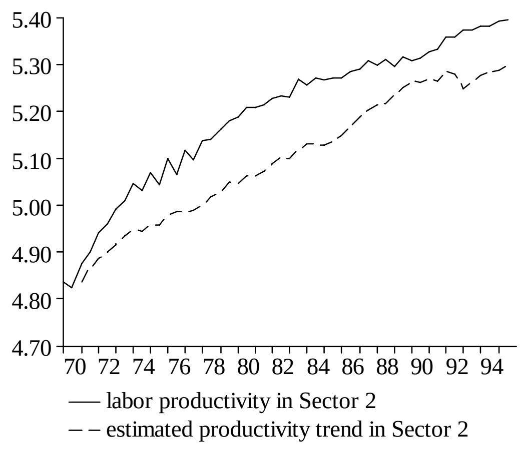


Fig. 6a Response in employment to a labor shock

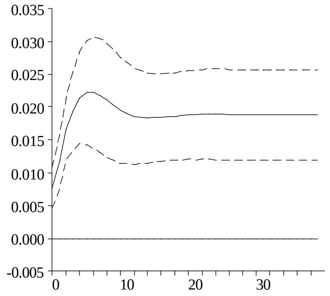


Fig. 7a Response in employment to a Sector 1 shock

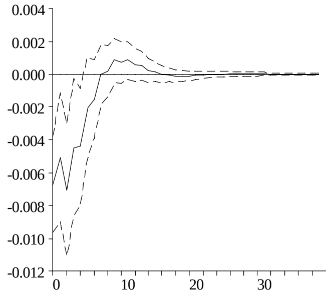


Fig. 8a Response in employment to a Sector 2 shock

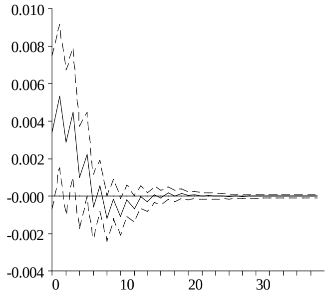


Fig. 6b Response in production in Sector 1 to a labor shock

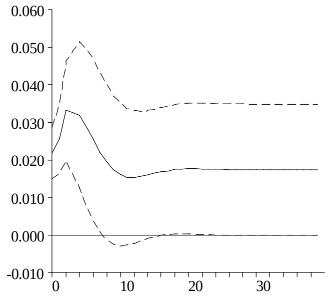


Fig. 7b Response in production in Sector 1 to a Sector 1shock

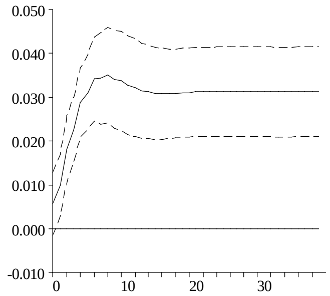


Fig. 8b Response in production in Sector 1 to a Sector 2 shock

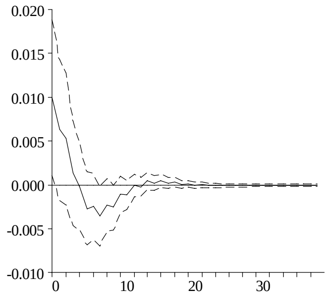


Fig. 6c Response in production in Sector 2 to a labor shock

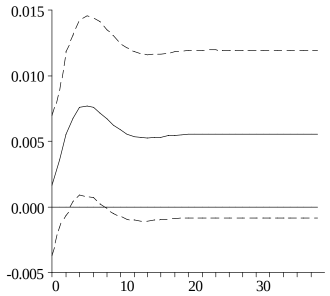


Fig. 7c Response in production in Sector 2 to a Sector 1 shock

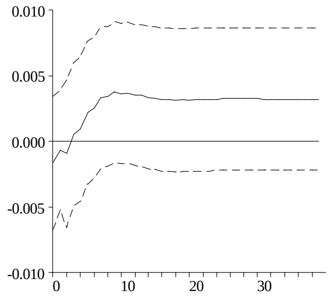
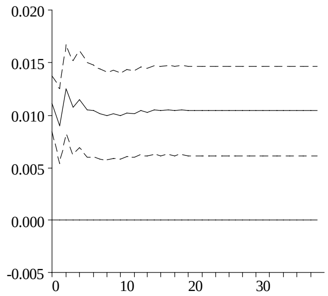


Fig. 8c Response in production in Sector 2 to a Sector 2 shock



Sammanfattning

I pappret undersöks hur produktionen i två sektorer är relaterade till varandra. Har chocker som uppstår i en sektor någon effekt på produktionen i den andra sektorn? Frågan är relevant både ur konjunkturhänseende och ur ett tillväxtperspektiv. En strukturell VAR används i analysen. Tre variabler, sysselsättningen i näringslivet, och förädlingsvärdet i industrin (Sektor 1) och i övrigt näringsliv (Sektor 2). Effekter av tre olika chocker med permanenta effekter studeras: en gemensam, tolkad som en chock till en stokastisk trend i arbetsutbudet, och två sektorsspecifika. Sysselsättningen antas följa arbetsutbudstrenden på lång sikt, och arbetsutbudstrenden tillåts påverka produktionen i båda sektorerna på lång sikt. Två alternativa identifierande antagande används. I det första fallet antas produktionen i sektor 1 opåverkad av chocker som uppstår i sektor 2, samtidigt som produktionen i sektor 2 tillåts påverkas av sektor 1-chocker. Det är därför möjligt att testa om sektor 1-chocker har permanenta effekter på produktionen i sektor 2. I den andra skattningen tillåts att sektor 2 chocker påverkar produktionen i sektor 1 och det går att testa om sektor 2-chocker har långsiktiga effekter på produktionen i sektor 1.

Hypotesen att sektors-specifika chocker påverkar produktionen i den andra sektorn finner inget stöd i data, varför vi kan dra slutsatsen att sektorsvis produktion drivs av egna chocker och av gemensamma chocker till båda sektorerna.

Eftersom chocker med enbart temporära effekter inte är med i analysen, kan detta medföra svårigheter att i praktiken skilja mellan utbuds och efterfrågestörningar. Det skulle därför vara av intresse att testa effekter av sektorsvisa störningar i en modell med både permanenta och tillfälliga effekter.