Empirical Modelling of Japan's Markup and Inflation, 1976-2000

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1 Objective

The objective of this paper is to model Japan's markup and inflation using historical time series data covering the last quarter of the 20th century.

It is demonstrated that a cointegrated vector autoregressive (VAR) model reduces to a parsimonious data-congruent dynamic system subject to economic interpretations.

Interactions between prices and wages have been of great interest for macro-economists, and theoretical models for prices and wages have played a key role in the development of modern macroeconomics.

Tobin (1972) presents a two-dimensional theoretical model encompassing a *markup* equation, and the model is successful in describing dynamic interdependent relationships of prices and wages.

Regarding applied work using time series data, Sargan (1964) is a seminal paper in that he gives congruent econometric representations of wages and prices in the UK based on an equilibrium correction approach, the approach intimately linked to a cointegration analysis.

A recent empirical analysis of price and wage time series data was performed by Bårdsen, Jansen and Nymoen (2003), Marcellino and Mizon (2001), *inter alia*.

Macroeconomic time series data often exhibit non-stationary behaviour, and thus need to be treated as integrated processes rather than stationary.

The concept of *cointegration* introduced by Granger (1981) therefore plays an important role in time series econometrics, and a cointegrated VAR model developed by Johansen (1988, 1996) has become a major econometric tool for macro and financial economists.

The cointegrated VAR analysis is well fitted in *general-to-specific* modelling approach, due to the fact that the analysis usually starts with the investigation of general unrestricted VAR models.

Hendry and Mizon (1993) discuss a model reduction procedure using the cointegrated VAR model. See also Hendry and Doornik (1994), Kurita (2007) for the general-to-specific modelling methodology using the cointegrated VAR analysis.

Furthermore, the concept of *weak exogeneity* introduced by Engle, Hendry and Richard (1983) also plays an important part in econometrics.

Weak exogeneity permits us to model a partial or conditional system alone, instead of a full system, for the purpose of making efficient statistical inferences on parameters of interest.

Weak exogeneity in the cointegrated VAR system is explored by Johansen (1992) and Urbain (1992).

Cointegration and weak exogeneity provide a methodological basis for empirical investigation pursued in this paper.

This paper aims to achieve a data-congruent parsimonious representation of markup and inflation using Japan's historical time series data.

The empirical exploration sheds useful light on deeper understanding of the Japanese economy in 1976 - 2000, a quarter-century period of Japan's economic turmoil, during which an asset-price bubble economy took place and then collapsed, leading to Japan's lost decade.

This paper adopts a cointegrated VAR approach to modelling the data of markup, inflation and various other macroeconomic series in Japan.

The analysis indicates the existence of a long-run economic linkage interpreted as an empirical representation of *countercyclical markup*, see Rotemberg and Woodford (1999), *inter alia*.

A set of variables in the cointegrated system, apart from markup and inflation, are judged to be weakly exogenous for parameters of interest, thereby allowing us to estimate a partial model given the weakly exogenous variables.

The model reduction is then conducted so as to achieve a parsimonious dynamic econometric system for markup and inflation.

It is noteworthy that such a stable structure has been revealed from the analysis of the data covering the period of Japan's economic turmoil.

2 Outline

1. Countercyclical Markup and Inflation Dynamics

2. Cointegrated VAR Model

3. Empirical Analysis of Japan's Time Series Data

4. Conclusion

3 Countercyclical Markup and Inflation Dynamics

This section introduces a set of economic variables to be analysed in a cointegrated VAR model.

Since the aim of this paper is to estimate an empirical model for Japan's markup and inflation, it is necessary to conceive a plausible long-run economic relationship associated with these two variables

— an interpretable economic linkage which may correspond to an empirical cointegrating relation estimated from the data.

To this end, let us suppose that markup pricing is formulated as

$$P_t = \theta_t \left(\frac{W_t}{A_t} \right), \tag{1}$$

where P_t is the price level, W_t is the nominal wage, A_t is the labour productivity, and θ_t is the markup. Let us introduce the output Y_t and define $y_t = \log Y_t$.

In order to map (1) to an observable relationship subject to an empirical investigation, the following specifications of A_t and θ_t are assumed:

$$\log \theta_t = -\delta \Delta y_t + u_t \quad \text{and} \quad \log A_t = \eta t + \upsilon_t, \tag{2}$$

where u_t and υ_t represent stationary error terms capturing unspecified dynamics.

The specification of θ_t is based on the fact that the markup tends to be moderately *countercyclical* with economic growth.

See Blanchard and Fisher (1989, Ch.9), Solon, Barsky and Parker (1994), Rotemberg and Woodford (1999), *inter alia*.

The productivity A_t is specified to be an exponential function of deterministic trend in (2), based on the assumption that productivity growth should be approximated to a stable upward trending path.

Substituting (2) into (1) and taking logs of both sides can lead to

$$p_t - (w_t - \eta t) + \delta \Delta y_t \sim stationary, \tag{3}$$

for $p_t = \log P_t$ and $w_t = \log W_t$. Equation (3) is a candidate for a cointegrating relationship based on the notion of countercyclical markup.

Let us define productivity-adjusted wage $w_t^* = w_t - \eta t$ for future reference. An inflation process should be dependent on the markup, possibly adjusting to disequilibrium errors represented by (3).

It is therefore reasonable to conceive the following bivariate equilibrium correction model (ECM) capturing countercyclical markup and inflation dynamics simultaneously:

$$\begin{aligned}\begin{bmatrix} \Delta\left(p_t - w_t^*\right) \\ \Delta^2 p_t \end{bmatrix} &= \begin{bmatrix} \zeta_1 \\ \zeta_2 \end{bmatrix} + \begin{bmatrix} \lambda_1 \\ \lambda_2 \end{bmatrix} \left[p_{t-1} - w_{t-1}^* + \delta \Delta y_{t-1}\right] \\ &+ \sum_{i=1}^{l-1} \begin{bmatrix} \gamma_{11,i} & \gamma_{12,i} \\ \gamma_{21,i} & \gamma_{22,i} \end{bmatrix} \begin{bmatrix} \Delta\left(p_{t-} - w_t^*\right) \\ \Delta^2 p_t \end{bmatrix} + \begin{bmatrix} \nu_{1,t} \\ \nu_{2,t} \end{bmatrix},\end{aligned} \tag{4}$$

where $\nu_{j,t}$ for $j = 1, 2$ is a set of stationary processes consisting of omitted short-run dynamic terms and innovations.

Such a bivariate system as (4) may be seen as an empirical representation of Tobin's wage-price model (Tobin, 1972).

This paper is interested in estimating an empirical dynamic model for markup and inflation, thus the ECM given by (4) is seen as a target representation which should be achieved as result of general-to-specific econometric modelling.

Furthermore, it is often pointed out that the spread between the long and short term interest rates contains information about expected future economic growth.

See Stock and Watson (1989), Bernard and Gerlach (1998), Hamilton and Kim (2002), and Ichiue (2004), *inter alia*.

The interest rate differential or yield spread, denoted by rs_t, may play a significant role in the short-run dynamics, or $\nu_{i,t}$, of the ECM.

4 Cointegrated VAR Model

The argument so far allows us to introduce an $I(1)$ cointegrated VAR(k) model encompassing (4), formulated as

$$X_t = (p_t - w_t, \Delta p_t, \Delta y_t, rs_t)', \tag{5}$$

and

$$\Delta X_t = \alpha\left(\beta', \gamma\right)\binom{X_{t-1}}{t} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \varepsilon_t, \quad \text{for} \quad t = 1, ..., T, \tag{6}$$

where a sequence of innovations ε_t has independent and identical normal $N(0, \Omega)$ distributions conditional on $X_{-k+1}, ..., X_0$, and $\alpha, \beta \in \mathbf{R}^{4\times r}$ for $r < 4$, $\gamma \in \mathbf{R}^{r\times 1}$, $\mu \in \mathbf{R}^{4\times 1}$ and $\Gamma_i \in \mathbf{R}^{4\times 4}$. Let $\beta^{*\prime} = (\beta', \gamma)$ and $X^*_{t-1} = \left(X'_{t-1}, t\right)'$ for future reference.

Johansen (1996) demonstrates details of likelihood-based inference for these parameters. In equation (6) α is referred to as adjustment vectors, while $\beta^{*\prime}$ is called cointegrating vectors.

In practice, there may be a case where both p_t and w_t exhibit $I(2)$-type non-stationary behaviour.

Considering the presence of a markup relation between p_t and w_t, it is reasonable to conjecture that $p_t - w_t$ is seen as an $I(1)$ process as a result of the removal of the common $I(2)$ stochastic trend or nominal-to-real transformation (see Kongsted, 2005).

In addition, there is a possibility that Δy_t may be seen as a stationary process rather than $I(1)$.

As shown in Johansen (1996, page 74), it is possible to include stationary variables in the VAR model as long as they are relevant in terms of economic theory and insight.

As the cointegrating rank r is unknown in practice, the rank needs to be determined based on the data analysis. A log-likelihood ratio ($\log LR$) test statistic consists of the null hypothesis of r cointegration rank $\mathsf{H}(r)$ against the alternative hypothesis $\mathsf{H}(p)$, and its asymptotic quantiles are provided by Johansen (1996, Ch.15).

Determining the cointegrating rank in the VAR model allows us to test various restrictions on α, β and γ in order to pursue the adjustment structure and cointegrating relationships subject to economic interpretation.

Cointegrating relationships, embodied by $\beta^{*\prime} X_{t-1}^{*}$, correspond to a set of stationary linear combinations, acting as equilibrium correction mechanisms in the VAR model.

Thus it is necessary to check if the theoretical relationship (3) belongs to a class of estimated cointegrating relationships.

Next, in order to derive the bivariate system (4) from the cointegrated VAR model (6) as a partial or conditional data-representation, let the process be decomposed as $X_t = \left(X_{1t}', X_{2t}'\right)'$ for $X_{1t} = (p_t - w_t, \Delta p_t)'$ and $X_{2t} = (\Delta y_t, rs_t)'$.

The parameters and error terms appearing in (6) are also expressed as

$$\alpha = \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix}, \Gamma_i = \begin{pmatrix} \Gamma_{1,i} \\ \Gamma_{2,i} \end{pmatrix}, \ \mu = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix}, \ \varepsilon_t = \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{pmatrix}, \Omega = \begin{pmatrix} \Omega_{11} & \Omega_{12} \\ \Omega_{21} & \Omega_{22} \end{pmatrix}.$$

We are interested in estimating (4), or a bivariate system for X_{1t} conditional on X_{2t} with no loss of information.

Suppose $\alpha_2 = 0$, then (6) is decomposed into a model for X_{1t} conditional on X_{2t} and a marginal model for X_{2t} as follows:

$$\Delta X_{1t} = \omega \Delta X_{2t} + \alpha_1 \beta^{*\prime} X^{*}_{t-1} + \sum_{i=1}^{k-1} \widetilde{\Gamma}_{1,i} \Delta X_{t-i} + \widetilde{\mu}_1 + \widetilde{\varepsilon}_{1,t}, \qquad (7)$$

$$\Delta X_{2t} = \sum_{i=1}^{k-1} \Gamma_{2,i} \Delta X_{t-i} + \mu_2 + \varepsilon_{2,t}, \qquad (8)$$

where

$$\omega = \Omega_{12}\Omega_{22}^{-1},\ \widetilde{\Gamma}_{1,i} = \Gamma_{1,i} - \omega \Gamma_{2,i},\ \widetilde{\mu}_1 = \mu_1 - \omega \mu_2,\ \widetilde{\varepsilon}_{1,t} = \varepsilon_{1,t} - \omega \varepsilon_{2,t},$$

and

$$\begin{pmatrix} \widetilde{\varepsilon}_{1,t} \\ \varepsilon_{2,t} \end{pmatrix} = N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \Omega_{11.2} & 0 \\ 0 & \Omega_{22} \end{pmatrix} \right],$$

for

$$\Omega_{11.2} = \Omega_{11} - \Omega_{12}\Omega_{22}^{-1}\Omega_{21}.$$

Note that $\beta^{*\prime} X^{*}_{t-1}$ is not embedded in the marginal model (8).

It is then possible to say that (4) may correspond to the conditional model (7).

Under the condition of $\alpha_2 = 0$, X_{2t} is said to be weakly exogenous for the following parameters of interest:

$$\alpha_1,\ \beta^*,\ \omega,\ \widetilde{\Gamma}_{1,1}, ..., \widetilde{\Gamma}_{1,k-1},\ \widetilde{\mu}_1,\ \text{and}\ \Omega_{11.2}. \tag{9}$$

The parameters in (9) correspond to those appearing in (4), and may therefore be treated as the set of parameters of interest.

As long as the condition for weak exogeneity $\alpha_2 = 0$ is satisfied, the parameters of interest (9) can be estimated from the conditional model (7) alone without loss of information, with no need for the estimation of the marginal model (8).

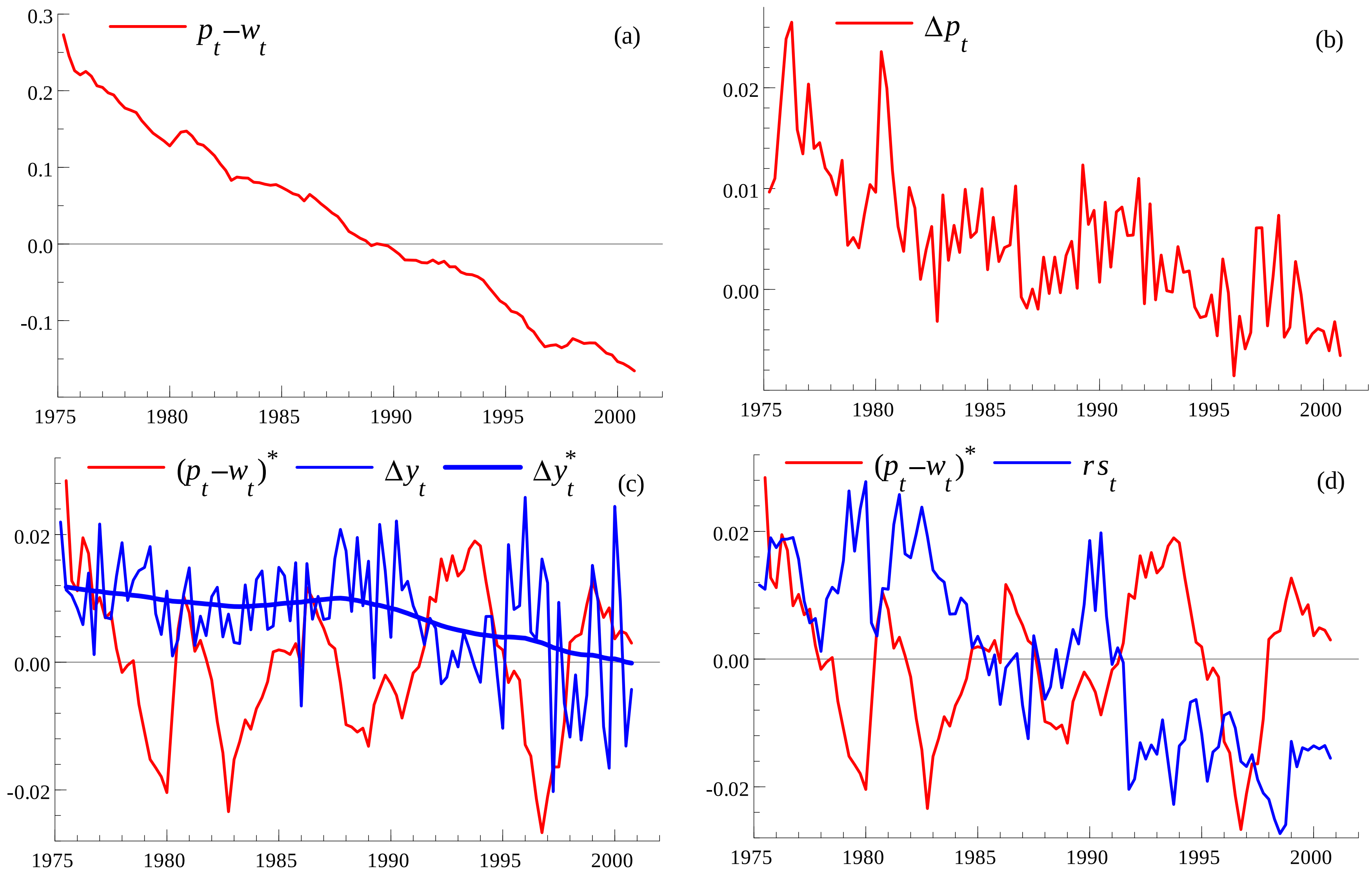
5 Empirical Analysis of Japan's Time Series Data

5.1 An Overview of the Data

An overview of Japan's time series data for

$$X_t = (p_t - w_t, \Delta p_t, \Delta y_t, rs_t)'$$

is presented below.



$p_t - w_t$
(a)
Δp_t
(b)
$(p_t - w_t)^*$
Δy_t
Δy_t^*
(c)
$(p_t - w_t)^*$
rs_t
(d)
0.3
0.2
0.1
0.0
-0.1
0.02
0.01
0.00
-0.02
1975
1980
1985
1990
1995
2000

5.2 Estimating the Unrestricted VAR Model

The empirical analysis commences with a general unrestricted VAR(5) model incorporating two deterministic terms, a constant and linear trend.

The inclusion of deterministic trend can be justified as it is treated as a local approximation to productivity; see the argument above.

F-tests for the lag order determination indicate that variables at lag length 5 seem to be irrelevant so the VAR(5) model reduces to the VAR(4) model.

Some evidence is then found for significance of a variable with lag length 4, suggesting that further model reduction is likely to be inappropriate.

The VAR(4) model is chosen for further analysis, so that the sample period effective for estimation is 1976.3 - 2000.4 and the number of observations is 98.

The unrestricted VAR model is a purely statistical representation, so the estimated coefficients are not necessarily subject to economic interpretation.

After identifying the cointegrating relations and conducting the model reduction, it is possible to pursue such interpretation.

The unrestricted VAR model provides a starting point towards a parsimonious representation, and should therefore pass a set of residual diagnostic tests such as non-autocorrelation and normality.

Single equation tests	$p_t - w_t$	Δp_t	Δy_t	rs_t
Autocorr. [F_{ar}(6,74)]	1.48 [0.20]	1.47 [0.20]	0.50 [0.81]	1.37 [0.24]
ARCH [F_{arch}(6,68)]	0.18 [0.98]	0.09 [1.00]	1.09 [0.38]	0.79 [0.58]
Hetero.[F_{het}(34,45)]	0.61 [0.93]	0.72 [0.84]	0.87 [0.67]	0.49 [0.98]
Normality [χ^2_{nd}(2)]	3.16 [0.21]	3.02 [0.22]	3.66 [0.16]	2.22 [0.33]

Vector tests			
Autocorr.[F_{ar}(16,223)]	1.22 [0.25]	Hetero.[F_{het}(340,374)]	0.66 [1.00]
- [F_{ar}(96,212)]	1.04 [0.41]	Normality [χ^2_{nd}(8)]	7.17 [0.52]

Note. Figures in brackets are p-values.

The diagnostic tests of the VAR(4) model are given in the table above.

Most of the test results are given in the form $F_j(k, T-l)$, which means an approximate F-test against the alternative hypothesis j :

kth-order serial correlation (F_{ar}: see Godfrey, 1978; Nielsen, 2007), kth-order autoregressive conditional heteroscedasticity (F_{arch}: see Engle, 1982), heteroscedasticity (F_{het}: see White, 1980), and a chi squared test for normality (χ^2_{nd}: see Doornik and Hansen, 1994).

The diagnostic test statistics are all insignificant, thereby allowing us to conclude that this model is subject to the subsequent likelihood-based cointegration analysis and model reduction.

5.3 Choosing the Cointegrating Rank

The table below presents two types of $\log LR$ test statistics for the choice of r, so-called trace test statistics ($trace$) and maximum eigenvalue test statistics ($max.eigen.$). The modulus of the six largest roots of a companion matrix for the VAR model are also provided in the table.

	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
$trace$	64.92[0.04]*	40.51[0.08]	23.77[0.09]	9.66[0.15]
$max.eigen.$	24.41[0.32]	16.73[0.48]	14.12[0.25]	9.66[0.15]

mod ($r = 1$)	1.00	1.00	1.00	0.78	0.78	0.76
mod ($r = 2$)	1.00	1.00	0.83	0.83	0.79	0.78

$Note.$ * denotes significance at the 5% level.

According to the first panel, the trace test rejects the null hypothesis of $r = 0$ and does not reject the remaining ones at the 5% significance level, hence supporting $r = 1$. In contrast, the maximum eigenvalue test does not reject any null hypotheses at the same level.

As discussed by Juselius (2006, Ch.8), the choice of the cointegrating rank is a very difficult task and so we should make use of as much additional information as possible in order to determine the rank.

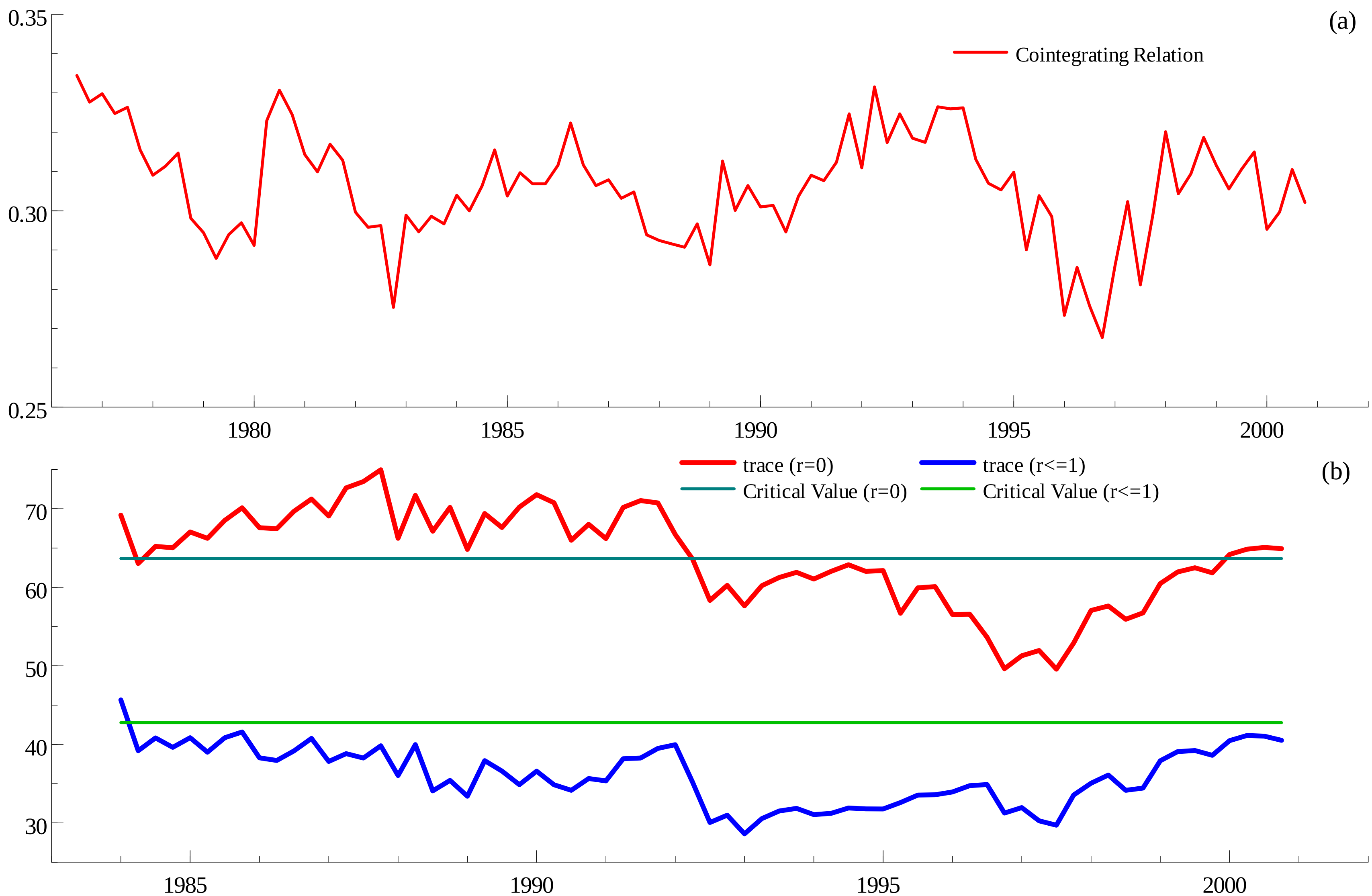
The second panel, motivated by the results of the trace test in the first panel, provides modulus (denoted **mod**) of the six largest eigenvalues of a companion matrix of the VAR model restricted with $r = 1$ or with $r = 2$.

No eigenvalue over 1.0 suggests that the model does not include any explosive root, and all the eigenvalues apart from the imposed unit roots are distinct from unity.

Judging from these outcomes, the restriction of $r = 1$ seems to be appropriate for the description of the data.

In order to consolidate the argument for $r = 1$, the figure below presents time series plots of the estimated cointegrating combination under the restriction of $r = 1$.

The figure displays recursive plots of some of the trace test statistics.



(a)
Cointegrating Relation
0.35
0.30
0.25
1980
1985
1990
1995
2000
(b)
trace (r=0)
trace (r<=1)
Critical Value (r=0)
Critical Value (r<=1)
70
60
50
40
30
1985
1990
1995
2000

5.4 Identifying the Long-Run Economic Relationship

This sub-section explores valid restrictions on the adjustment and cointegrating space in the $I(1)$ cointegrated VAR system. The determination of the cointegrating rank, or $r = 1$, enables us to inspect such restrictions using a standard χ^2-based asymptotic inference.

Weak exogeneity plays an important role in the model reduction procedure. Testing weak exogeneity in the $I(1)$ cointegrated system corresponds to checking zero restrictions on elements of α.

If rs_t and Δy_t are judged to be weakly exogenous for parameters of interest, one has only to model the conditional system for the variables $p_t - w_t$ and Δp_t in order to conduct inferences with no loss of information.

With regard to restrictions on β, it is of interest to identify the long-run empirical relationship interpreted as the markup equation given by (3).

It is expected that the inflation process could play a little or no role in the long-run relationship, so the exclusion of Δp_t from the long-run cointegrating relation should be examined.

Thus, the following restrictions on α and β^* are jointly investigated:

1. zero restrictions on the elements of α corresponding to rs_t and Δy_t,

2. a zero restriction on the element of β^* corresponding to Δp_t.

A set of restricted estimates is given in the table below, together with the corresponding log LR test statistic and p-value.

	$p_t - w_t$	Δp_t	Δy_t	rs_t	t	$\chi^2(3)$
$\widehat{\alpha}'$	-0.19	0.19	0	0	$-$	4.76[0.19]
	(0.05)	(0.05)	$(-)$	$(-)$		
$\widehat{\beta}^{*\prime}$	1	0	1.12	0.002	0.00389	
	$(-)$	$(-)$	(0.51)	(0.004)	(1.51e-04)	

Note. The figures in the parentheses are standard errors.

The table shows that the set of hypotheses is not rejected at the 5% level, indicating that rs_t and Δy_t are weakly exogenous for the parameters of interest and Δp_t can be excluded from the cointegrating space.

It turns out that the coefficient for rs_t in the cointegrating space is insignificant and that for Δy_t is close to unity. Thus, it would be worthwhile to test additional restrictions as follows:

1. a zero restriction on the element of β^* corresponding to rs_t,

2. a unitary restriction on the element of β^* corresponding to Δy_t.

The table below presents a set of restricted estimates, together with the corresponding log LR test statistic and p-value.

	$p_t - w_t$	Δp_t	Δy_t	rs_t	t	$\chi^2(5)$
$\widehat{\alpha}'$	-0.18 (0.04)	0.19 (0.04)	0 $(-)$	0 $(-)$	$-$	5.01[0.41]
$\widehat{\beta}^{*\prime}$	1 $(-)$	0 $(-)$	1 $(-)$	0 $(-)$	0.00381 (7.15e-05)	

Note. The figures in the parentheses are standard errors.

Again, the set of hypotheses is not rejected at the 5% level, indicating that the long-run economic relationship is given by

$$p_t - (w_t - 0.00381t) + \Delta y_t. \tag{10}$$

Interpreting the linear trend as an approximation of labour productivity growth, this cointegrating relation indicates that a markup over productivity-adjusted wages tends to move in the opposite direction to the real output growth.

Relationship (10) is therefore interpreted as (3), consistent with the accepted view of *countercyclical markup*, discussed in the literature of macro and labour economics.

See Blanchard and Fisher, 1989, Ch.9; Solon, Barsky and Parker, 1994; Rotemberg and Woodford, 1999; Romer, 2001, Ch.5, *inter alia*.

Various theories have been developed to explain this pattern of countercyclical markup such as collusion in imperfect competition and kinked demand curves, see the above references.

Under these economic interpretations, (10) can be justified as the representation of a meaningful long-run economic relationship.

5.5 A Parsimonious Equilibrium Correction Model

We are now in a position to achieve an equilibrium correction model for Japan's markup and inflation, corresponding to the bivariate system (4) above.

The concept of productivity-adjusted wages is used to provide appropriate empirical markup, that is, an adjusted wage index is defined as $w_t^* = w_t - 0.00381t$ such that markup is represented by $p_t - w_t^*$.

As a result of this adjustment, the sample mean of $\Delta\left(p_t - w_t^*\right)$ appears to be around zero rather than a negative value.

The starting point of the analysis is to map the data to the $I(0)$ space by differencing and using the restricted cointegrating combination.

We then estimate a two-dimensional $I(0)$ VAR system for $\Delta\left(p_t - w_t^*\right)$ and $\Delta^2 p_t$ conditional on Δrs_t and $\Delta^2 y_t$.

A set of insignificant terms, $\Delta^2 p_{t-2}$, $\Delta^2 p_{t-3}$, $\Delta^2 y_{t-1}$, $\Delta^2 y_{t-2}$ and Δrs_{t-3}, are dropped so as to reach a parsimonious VAR model.

Short-run dynamics with large standard errors then continue to be removed, that is, $\Delta\left(p_{t-2} - w_{t-2}^*\right)$ from the equation for $\Delta\left(p_t - w_t^*\right)$, and Δrs_{t-1} and Δrs_{t-2} from the $\Delta^2 p_t$ equation.

Imposing constraints on some of the coefficients which have similar size in order to seek a parsimonious representation, an empirical price-wage mechanism is attained as follows:

$$\Delta\left(\widehat{p_t - w_t^*}\right) = \underset{(0.05)}{-0.12}\,\Delta^a rs_t - \underset{(0.02)}{0.08}\,\Delta^{2a} y_t + \underset{(0.09)}{0.27}\,\Delta(p_{t-1} - w_{t-1}^*)$$
$$- \underset{(0.08)}{0.14}\,\Delta^2 p_{t-1} - \underset{(0.06)}{0.19}\,\Delta rs_{t-1} + \underset{(0.07)}{0.11}\,\Delta\left(p_{t-3} - w_{t-3}^*\right)$$
$$- \underset{(0.03)}{0.17}\,ecm_{t-1} + \underset{(0.01)}{0.05},$$
$$\widehat{\sigma} = 0.0038,\ F_{ar}(6{,}81) = 1.97[0.08],$$
$$\chi^2_{nd}(2) = 2.75[0.25],\ F_{arch}(6{,}78) = 0.35[0.91],\ F_{het}(20{,}69) = 1.02[0.46],$$

$$\Delta^2\widehat{p}_t = - \underset{(0.06)}{0.12}\,\Delta rs_t - \underset{(0.03)}{0.09}\,\Delta^2 y_t - \underset{(0.09)}{0.64}\,\Delta^2 p_{t-1} - \underset{(0.07)}{0.17}\,\Delta^a(p_{t-1} - w_{t-1}^*)$$
$$+ \underset{(0.08)}{0.4}\,\Delta(p_{t-2} - w_{t-2}^*) - \underset{(0.03)}{0.16}\,ecm_{t-1} + \underset{(0.01)}{0.05},$$
$$\widehat{\sigma} = 0.0038,\ F_{ar}(6{,}81) = 2.18[0.053],$$
$$\chi^2_{nd}(2) = 2.73[0.26],\ F_{arch}(6{,}78) = 0.33[0.92],\ F_{het}(20{,}69) = 1.02[0.45],$$

$$Vector\ Tests : F_{ar}(24{,}154) = 1.18[0.27],\ \chi^2_{nd}(4) = 4.69[0.32],\ F_{het}(60{,}200) = 0.90[0.68],$$

where

$$ecm_t = p_t - w_t^* + \Delta y_t,$$

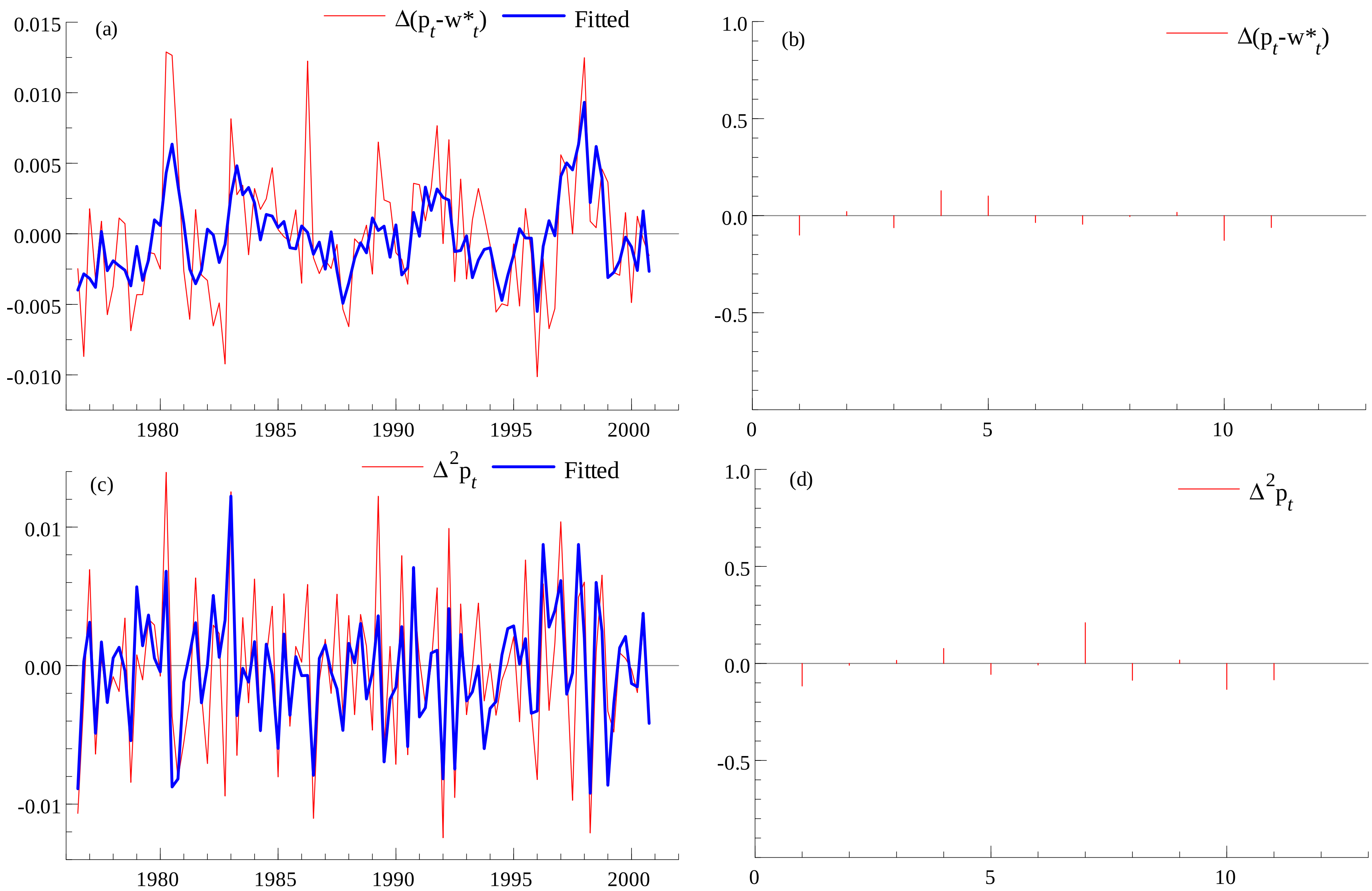
and

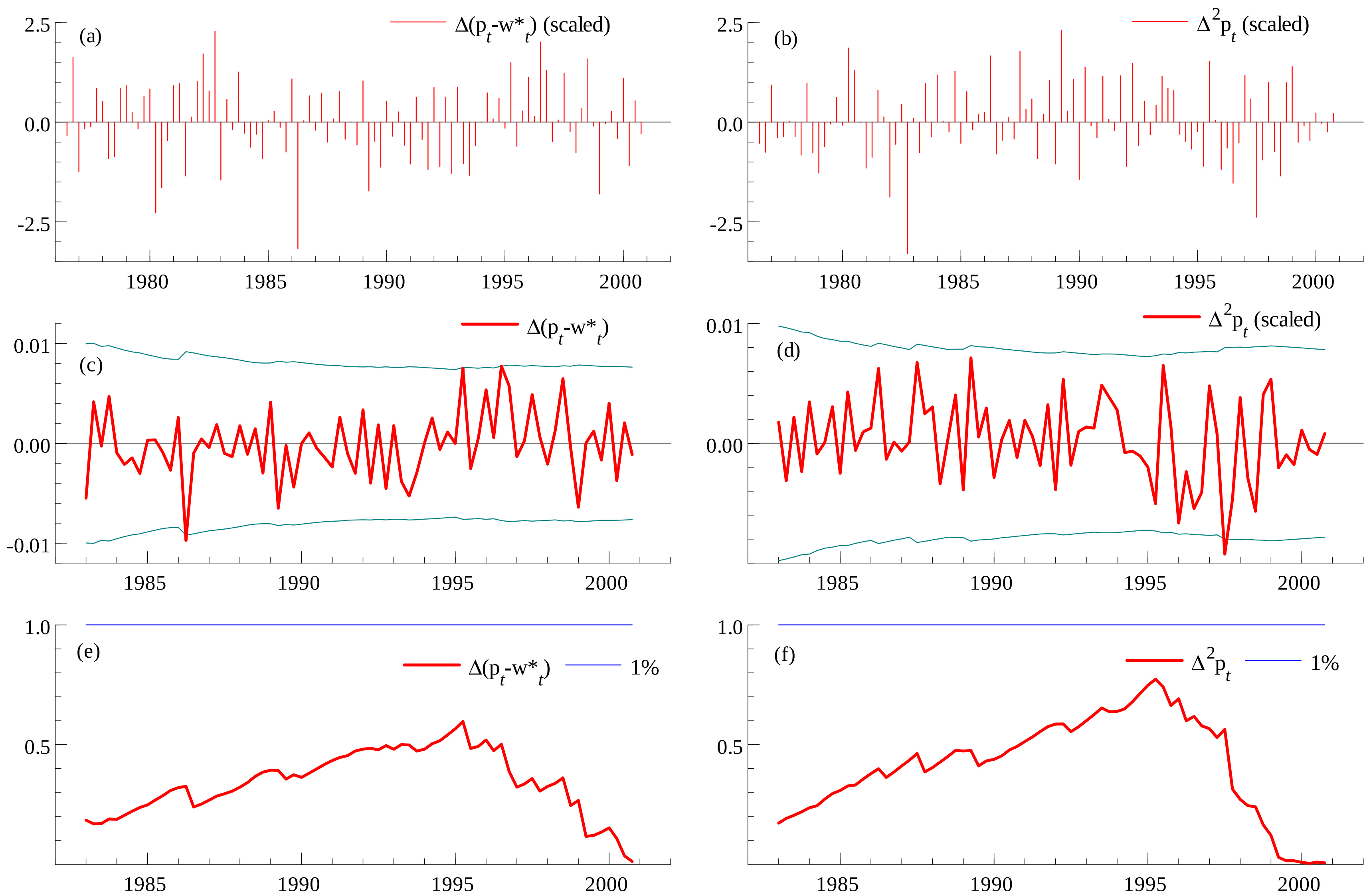
$$\Delta^a rs_t = \Delta rs_t + \Delta rs_{t-2},\ \Delta^{2a} y_t = \Delta^2 y_t + \Delta^2 y_{t-3},$$
$$\Delta^a (p_{t-1} - w_{t-1}^*) = \Delta(p_{t-1} - w_{t-1}^*) - \Delta(p_{t-3} - w_{t-3}^*).$$

The method of constrained full-information maximum likelihood was used to estimate the model above.

The standard errors on coefficients are given in parentheses.

The test statistics for a single equation are reported under each equation, and the statistics for the whole system (the vector tests) are reported under the two equations.


(a)
Δ(p_t-w*_t)
Fitted
0.015
0.010
0.005
0.000
-0.005
-0.010
1980
1985
1990
1995
2000
(b)
Δ(p_t-w*_t)
1.0
0.5
0.0
-0.5
0
5
10
(c)
Δ^2p_t
Fitted
0.01
0.00
-0.01
1980
1985
1990
1995
2000
(d)
Δ^2p_t
1.0
0.5
0.0
-0.5
0
5
10



(a)
Δ(p_t-w*_t) (scaled)
2.5
0.0
-2.5
1980
1985
1990
1995
2000
(b)
Δ²p_t (scaled)
2.5
0.0
-2.5
1980
1985
1990
1995
2000
(c)
Δ(p_t-w*_t)
0.01
0.00
-0.01
1985
1990
1995
2000
(d)
Δ²p_t (scaled)
0.01
0.00
1985
1990
1995
2000
(e)
Δ(p_t-w*_t)
1%
1.0
0.5
1985
1990
1995
2000
(f)
Δ²p_t
1%
1.0
0.5
1985
1990
1995
2000

None of the diagnostic test statistics are significant, suggesting that the parsimonious system is a data-congruent representation.

Next, let us consider interpretation of the parsimonious model.

A change in the interest rate differential has a negative effect on the markup and inflation growth; this may reflect information on expected economic growth and inflation contained in the differential.

In line with the interest rate differential, an acceleration of the real output growth also has a negative influence on the markup and inflation growth.

This is consistent with the countercyclical behaviour found in the cointegrating relation, and could be interpreted as a short-run reflection of this.

6 Summary and Conclusion

This paper, using a cointegrated VAR methodology, estimates a data-congruent econometric model for Japan's markup and inflation.

The analysis provides evidence for the presence of a long-run economic relationship in the data, interpreted as an empirical representation of countercyclical markup.

A set of variables in the cointegrated system except markup and inflation are judged to be weakly exogenous for parameters of interest, thereby enabling us to estimate a partial system given the weakly exogenous variables.

The model reduction is then conducted in order to attain a parsimonious dynamic econometric system.

The preferred parsimonious system has passed a battery of diagnostic tests, thereby being judged to be a data-congruent representation of countercyclical markup and inflation dynamics.

It should be noted that such a stable model has been estimated from the analysis of the data covering the period of Japan's economic turmoil.

The empirical exploration sheds useful light on deeper understanding of the Japanese economy in the last quarter of the 20th century.