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THE VARIANCE HYPOTHESIS ON OUTPUT-INFLATION TRADEOFF - EVIDENCE FROM SCANDINAVIA***

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Abstract

This paper contains some empirical tests for the variance hypothesis according to which the output-inflation tradeoff varies inversely with the variance of nominal aggregate demand. Both Finnish quarterly and annual cross-country data on Scandinavian countries are used. Finnish time series data lies in conformity with the hypothesis, whereas the tests with cross-country data, particularly when Zellner's seemingly unrelated regression estimation procedure is used, fail to point to the supporting evidence for the variance hypothesis.

1. INTRODUCTION

The Phillips-curve hypothesis, according to which the output-inflation tradeoff results from the structural features of the economy and is independent of policies being pursued, was challenged in late sixties by the 'natural' rate of unemployment hypothesis (NRH). It suggested that attempts to move along the Phillips curve to decrease unemployment will be frustrated in the long run, but not in the short run, by changes in expectations that shift the curve. According to NRH, however, the effectiveness of activist Keyneasian aggregate demand policies rested on fooling people.[1)]

The rational expectations hypothesis (REH) provided an innovative approach to economic modelling by suggesting that expectations, instead of being arbitrary, are 'rational' in the sense of using information efficiently. If the natural rate of output is exogenous and prices always move to clear markets, then coupling the linear aggregate supply function with REH makes it possible to draw two strong conclusions: (i) deviations of output from its natural rate are pure white noise - there is no business cycle, and (ii) choice among aggregative monetary and fiscal policy rules is irrelevant for the stochastic behaviour of output and unemployment.[2)]

Modifications of the basic model to allow for serial correlation of output and thereby for business cycles do not necessarily invalidate the conclusion (ii) (see, e.g. Blinder and Fischer [5], Lucas [16] and Sargent [30]). This neutrality result does break down when any of the following assumptions is rejected: linearity of the aggregate supply function (Persson [29]), exogeneity of the natural rate of output (Fair [12] and assumption of the market-clearing (for a survey, see McCallum [23]).[3)]

Anyway, regardless of a priori notions on the plausibility of rational expectations models, they can be subjected to empirical tests. Although formal econometric evidence does not suggest their out-of-hand rejection (see, e.g. Barro [3], [4], and Small [32]), the evidence cannot be regarded as compelling because REH has not generally been tested in isolation, but in the context of some other hypotheses. In fact Sargent [31] has shown that it is literally impossible to distingquish between Keynesian and classical macroeconomic structures using only parameter estimates from a single policy regime.

Since the neutrality property alone places no restrictions on time-series data taken from a single policy regime, one way to test NRH cum REH is to find periods across which the policy regimes differ and to test the invariance of alternative models across regimes. Along these lines the use of cross-equation restrictions to test a natural rate-rational expectations model was first proposed by Lucas [18].[4)] Specifically, he argued that the higher is the variance of monetary and fiscal behaviour, the smaller are the effects of unanticipated monetary and fiscal actions on output and vice versa. According to this hypothesis the output-inflation tradeoff varies inversely with the variance of monetary and fiscal innovations. Lucas tested it for 18 countries over the period 1953-67 and found some support to it. The major evidence, however, rested only on two Latin American countries.

This paper contains further tests of this variance hypothesis on the output-inflation tradeoff in two respects: First, we use different sample, namely Finnish time-series quarterly data over the period 1948-77 and annual cross-country data on Scandinavian countries over the period 1950-76. Second, Lucas estimated equations one-by-one by using OLS.

But it might well be the case that the error terms across countries are contemporaneously correlated, in which case Zellner's seemingly unrelated equation estimation technique is the appropriate estimation procedure. Checking whether the results are sensitive to the use of single-equation estimation technique is our second aim.

We proceed as follows. Section II outlines the variance hypothesis by Lucas. Empirical results are presented in section III and finally, some conclusions follow.

II. THE VARIANCE HYPOTHESIS ON THE OUTPUT-INFLATION TRADEOFF

This section outlines the argument behind the variance hypothesis. All variables are expressed in log terms for convenience (for a more detailed treatment, see Lucas [18], Sargent [30], Ch XIII).

The term 'Phillips curve' has been used to refer to a positive association between output and inflation, which contradicts with any model, in which agents' decisions about real economic variables are homogenous of degree zero in nominal variables. Lucas' theory of 'Phillips curve' reconciles this empirical observation with NRH according to which permanent changes in the inflation rate will not alter the average output and unemployment.

The basic idea is to develop an operational model of "money illusion", in which real economic decisions depend on relative prices, but agents do not have sufficient information to distinguish perfectly between relative and general price movements. Under these circumstances inferences on relevant, but unobserved prices are made 'optimally' taking account of the stochastic nature of the economy. Moreover, aggregate price-quantity observations are viewed as intersection points of aggregate demand and aggregate supply in the spirit of new classical macroeconomics.

In an economy in which all trading occurs in a single competitive market there is "too much" information in the hands of economic agents for them ever to be "fooled" into altering real decision variables. Assume that suppliers are distributed over a large number of competitive markets with unevenly distributed demand. Suppliers know the current local price $P_t(z)$ in market z at time t, and the history of the economy, but not the current general price level P_t. In each market z the supply is a product of a long run trend component, common to all markets, and a cyclical component so that $y_t(z) = y_{nt} + y_{ct}(z)$. The long run component is assumed to be exogenous, while the cyclical component is determined by the observed local price $P_t(z)$ relative to the unobserved general price level P_t expected on the basis of information $I_t(z)$ available in z at time t, and by its own lagged value so that

$$(1) \qquad y_{ct}(z) = e(P_t(z) - E(P_t|I_t(z)) + my_{c,t-1}(z)$$

(for rationalizations of including $y_{c,t-1}(z)$, see [30], pp. 330-31).

The information set $I_t(z)$ relevant for estimation of the unobserved P_t consists of the observed local price $P_t(z) = P_t+z$ and of the history summarized in the mean of the current price level at time t, $\bar{P}_t$. Assuming that P_t and z are normally and independently distributed[5)] with $(\bar{P},s^2)$ and $(0,q^2)$ it can be shown that

$$(2) \qquad E(P_t|I_t(z)) = (1-w)P_t(z) + w\bar{P}_t$$

where $w = q^2/(q^2+s^2)$ = the fraction of the conditional variance in $P_t(z)$ which is due to relative price variation. Thus the higher is w, the smaller weight is placed on $P_t(z)$ in forming estimates of the unobserved general price level and vice versa. Combining $y_t(z)$, (1) and (2) and averaging over all markets gives the aggregate supply function

$$(3) \qquad y_{ct} = we(P_t-\bar{P}_t) + my_{c,t-1}$$

where the slope depends on the relative price parameter e and the signal extraction parameter w.

The equation (3) has a natural interpretation. The higher is the signal extraction parameter w, the more likely a change in the observed price reflects a relative rather than a general price change. If GNP has shown small variability, then suppliers tend to regard shocks as being specific to them rather than economy-wide and attribute most demand shocks to micro situation. In the extreme case of $s^2 = 0$ all shocks will be attributed to micro situation (w=1) and the slope of (3) reflects only the relative price parameter e. But when the variability of GNP increases, then agents begin to regard more of the shocks as being economy-wide rather than specific to them. In the other extreme case of $q^2=0$ all shocks will be attributed to macro situation (w=0) and the aggregate supply function (3) becomes vertical.

This dependence of the slope of the aggregate supply function on the ratio of variances of relative to general price movements has an important implication: the aggregate supply function is not predicted to remain unchanged across different aggregate demand regimes. E.g. an attempt by the authorities to try to exploit the tradeoff more fully by changing aggregate demand can be expected to increase s^2 relative to q^2 and thereby decrease the tradeoff. To see this and to complete the model Lucas postulates a demand function of the form $x_t = y_t + P_t$, where x_t = nominal aggregate demand, an exogenous shift variable. Assuming that dx_t is a sequence of independent, normal variates with mean u_x and variance s_x^2 and combining (3) with x_t and the assumption of rational price expectations gives the following equations for the equilibrium values of real output and the inflation rate (see Lucas [18], pp. 328-29).6)

(4) $$y_{ct} = -ru_x + r\Delta x_t + my_{c,t-1}$$

(5) $$\Delta P_t = -b + (1-r)\Delta x_t + r\Delta x_{t-1} - m\Delta y_{c,t-1}$$

where $r = we/(1+we)$ and b = the coefficient from $y_{nt} = a + bt$.

Earlier discussion indicates that the coefficient r varies from $r = e/(1+e)$, when $w=1$, to $r=0$, when $w=0$. According to the variance hypothesis the coefficient r is inversely ralated to the variability of nominal aggregate demand.[7)] Before turning to empirical tests it should be stressed that (4) and (5) do not consist of a two-equation model as claimed in Lucas [18], but they are transforms of each other (see Lucas [19]).

III. EMPIRICAL TESTS OF THE VARIANCE HYPOTHESIS

The main test is to fit equations (4) and (5) and compare the estimates of r with the variances of Δx_t. Some attention is also paid to the explanatory power of these equations.

III.1. Tests with Finnish quarterly data

Our first test was performed with Finnish quarterly data over the period 1948(I) - 1977(IV), which was divided into three subperiods with different variances of Δx_t. Table 1 gives some descriptive statistics.[8)]

Table 1. Descriptive statistics (Finland)

period	$\overline{\Delta y}_t$	$\overline{\Delta P}_t$	$var(y_{c,t})$	$var(\Delta P_t)$	$var(\Delta x_t)$
1948-1956	.0554	.0764	.0018	.0116	.0169
1957-1966	.0523	.0324	.0012	.0017	.0026
1967-1977	.0424	.0976	.0028	.0038	.0031

Data source: Finnish National Accounts

But how do we know that these subperiods correspond to different policy regimes? It might be the same regime with only different realizations of the underlying stochastic process. Testing for the equality of sample variances of Δx_t for subperiods produced the following values of F-statistic: F(I, II) = 6.50, F(I, III) = 5.45 and F(II, III) = 1.19, so that the first two are significantly different from each other at the 1 per cent level, while the last one is not even at the 5 per cent level. Nevertheless we want also to keep periods II and III separate for checking purposes.

The performance of equations (4) and (5) in accounting for movements in real output and inflation rate is summarized in Table 2, where the estimates for the period 1957-1977 is also reported. Subscripts y and p refer to (4) and (5) respectively. Numbers in parentheses are values of t-statistics and the first-order serial correlation coefficient for equation (4) has been calculated by Hildreth-Lu procedure. Estimates are OLS estimates for real output equation, while inflation rate equation has been estimated by restricting the sum of coefficients of Δx_t and Δx_{t-1} to unity.

Table 2. Periodic estimates of output and inflation rate equation

period	r^*_y	$\bar{R}^2_y$	ρ_y	r^*_p	$\bar{R}^2_p$	$D\text{-}W_p$
I 49-56	.1436 (2.1)	.145	-.04	.3517 (2.4)	.776	2.11
II 57-66	.1726 (1.6)	.644	.66	.6445 (1.1)	.481	1.76
III 67-77	.3011 (1.5)	.402	.84	.5693 (4.4)	.810	2.42
IV 57-77	.2260 (3.2)	.226	.62	.5697 (7.5)	.795	2.06

Critical t-values: $t_{.05,30} = 1.697$, $t_{.01,30} = 2.457$.

As far as the performance of the inflation rate equation for subperiods I, II and III is concerned, the estimates of r are inversely related to var(Δx_t) in full conformity with the variance hypothesis, while the evidence on real output equation is not totally straightforward. The goodness-of-fit statistics, however, are in line with the hypothesis. As concerns the results for subperiods I and IV the estimate of r is smaller for period I than for period IV in both equations, quite in line with the variance hypothesis.[10] Thus the periodic time series data from Finland supports the variance hypothesis.

III.2. Tests with Scandinavian cross-country data

Turn now to tests with annual cross-country data from five Scandinavian countries for which some descriptive statistics over the period 1950-1976 are given in Table 3.[11]

Table 3. Descriptive statistics (Scandinavia)

country	$\overline{\Delta y}_t$	$\overline{\Delta P}_t$	var($y_{c,t}$)	var(ΔP_t)	var(Δx_t)
Denmark	.0407	.0551	.00123	.00073	.00079
Finland	.0472	.0607	.00093	.00401	.00485
Iceland	.0493	.1351	.00742	.00761	.00742
Norway	.0430	.0459	.00090	.00104	.00133
Sweden	.0386	.0485	.00051	.00159	.00108

Data source: Yearbook of National Account Statistics

In estimating the equations (4) and (5) for Scandinavian cross-country data we proceeded as follows: The output equation was first estimated separately for each country by using OLS. Since Durbin's m-statistic revealed the existence of significant serial correlation, a correction for it was made by using Hildreth-Lu and Cochrane-Orcutt procedures. It turned out, however, that the error terms across countries were contemporaneously correlated, and the single equations did not involve the same explanatory variables. Since under these conditions the single equation estimation methods are inefficient, we also used Zellner's seemingly unrelated equation estimation technique (Zellner [36]). Finally, the inflation rate equation (5) was estimated by restricting the sum of coefficients of Δx_t and Δx_{t-1} to unity without applying, however, system estimation methods in this connection.

Start with the real output equation. The single equation estimates are reported in Table 4. After correcting for first-order serial correlation by Hildreth-Lu and Cochrane-Orcutt procedures Finland and Iceland - which are characterized by the highest $\mathrm{var}(\Delta x_t)$ - seem to have the lowest estimate of r, while Denmark - which is characterized by the lowest $\mathrm{var}(\Delta x_t)$ - seems to have the highest estimate of r. The goodness-of-fit statistics for these countries are also roughly in line with the idea that the real output equation should have poor fit, if GNP has shown high variability. On the other hand, estimates of r for Norway and Sweden are lower than that for Finland, while the variance hypothesis would predict the reverse. Moreover, the goodness-of-fit statistic for Norway is rather different from what one would expect on the basis of $\mathrm{var}(\Delta x_t)$ and the estimate of r for Sweden appears to be particularly sensitive for the first-order serial correlation correction.[12)]

Table 4. Single equation (OLS) estimates of output equation

country	r^*	$\bar{R}^2/F$	m	r^*	$\bar{R}^2/F$	ρ	r^*	$\bar{R}^2/F$	ρ
	$\rho = 0$			Hildreth-Lu			Cochrane-Orcutt		
Denmark	.4338	.649	.814	.7041	.787	.797	.7111	.461	.811
	(3.0)	23.27	(3.3)	(5.3)	45.41		(3.3)	11.27	(7.0)
Finland	.2364	.296	1.352	.4002	.432	.605	.4478	.124	.592
	(2.1)	6.03	(5.2)	(3.2)	10.11		(2.7)	2.71	(3.7)
Iceland	.1172	.349	.944	.3696	.476	.705	.1761	.254	.503
	(1.0)	7.46	(3.7)	(2.6)	11.90		(1.1)	5.08	(2.9)
Norway	.2050	.844	-.050	.2050	.844	-.054	.2050	.844	-.051
	(3.1)	65.92	(0.2)	(3.2)	65.50		(3.2)	65.80	(0.3)
Sweden	-.0950	.572	.572	.3002	.672	.900	.3510	.455	.997
	(0.7)	17.05	(2.1)	(2.0)	25.67		(0.1)	10.99	(58.6)

The period of estimation: 1952-1976, m indicates Durbin's m-statistic.

Critical t-values: $t_{.05,22} = 1.717$ $t_{.01,22} = 2.508$

Critical F-values: $F_{.05,2,22} = 3.44$ $F_{.01,2,22} = 5.72$

Concludingly, the single equation estimates for the real output equation, with Scandinavian cross-country annual data give mixed evidence in terms of the variance hypothesis.

If the error terms are contemporaneously correlated while the explanatory variables are not identical, then Zellner's seemingly unrelated regression (SUR) technique would be the appropriate estimation procedure. Compared with single equation methods, the gain in efficiency is 'large', when contemporaneous disturbances across equations are, while explanatory variables of various equations are not 'highly' correlated.[13)] The information on these matters is reported in Table 5. Correlations between explanatory variables are always significant at the 5 per cent level, which does not seem to be the case between error terms across countries.

Still, error terms display significant contemporaneous correlation in so many cases that it is of importance to check whether the results are sensitive to the use of single-equation estimation technique. In implementing SUR-technique both maximum likelihood and Zellner's two-stage Aitken estimators were used.[14)].

Results by applying the maximum likelihood technique are reported in Table 6. No correction for serial correlation has been made in the lack of computer program. The first column presents estimates of r when the covariance matrix of disturbances is diagonal so that the contemporaneous correlation of disturbances is ignored. The second column reports the maximum likelihood estimates and the last column deals with the case in which estimates of r were restricted to be equal across countries as a sort of the weak test of the variance hypothesis.

Table 5. Coefficients of correlation w.r.t. μ_{it} and Δx_{it}

country	Denmark	Finland	Iceland	Norway	Sweden
Denmark	1.0000	.4459	.5186	.5483	.5265
Finland	.4146	1.0000	.5867	.6093	.6708
Iceland	.2497	.2343	1.0000	.5392	.6919
Norway	.3518	.1750	.1834	1.0000	.8431
Sweden	.3920	.4267	.1508	.1947	1.0000

The coefficients of correlation w.r.t. the OLS residuals, μ_{it}, are reported on the (down) left, and the ones w.r.t. Δx_{it} on the (up) right. Critical values (according to t-test) are: .337 at .05 and .463 at .01 level.

Testing for the significance of the constraint $\sigma_{ij} = 0$ for $i \neq j$ the likelihood ratio test can be used. The value of χ^2-statistics is $\chi^2_{10} = 112.55$ whereas the critical value of χ^2-statistics is $\chi^2_{.01,10} = 23.21$. Thus accounting for contemporaneous correlation of error terms makes the performance of the real output equation significantly better. Consider next the hypothesis that r is equal across Scandinavian countries. The value of χ^2-statistic is $\chi^2_4 = 6.06$ whereas the critical value at the 10 per cent significance level is $\chi^2_{.10,4} = 7.78$. Thus the hypothesis that r is equal across Scandinavian countries cannot be rejected. Clearly, the maximum likelihood estimates fail to support the variance hypothesis.[15)]

Table 6. Maximum likelihood estimates of output equation

country	$r^*(\sigma_{ij}=0,\ i\neq j)$	r^*	$\bar{r}^*$
Denmark	.4477 (2.2)	.5736 (7.4)	.3780 (8.4)
Finland	.1818 (1.0)	.4424 (5.1)	.2780 (8.4)
Iceland	.1079 (0.6)	.3593 (3.8)	.3780 (8.4)
Norway	.2054 (1.6)	.3183 (5.6)	.3780 (8.4)
Sweden	-.0705 (0.2)	.3803 (3.6)	.3780 (8.4)

Numbers in parentheses are (asymptotic) t-values.

Turn now to results obtained by Zellner's two-stage Aitken estimation technique, which are reported in Table 7. Equations were estimated both with and without serial correlation correction and estimates in the former case were obtained by using Parks' procedure in accounting for serial correlation (see Parks [27].[16]

Using Buse's generalized goodness-of-fit measure, R_G^2, (see Buse [7]) the system of output equations turns out to perform better when contemporaneous and serial correlation of disturbances are accounted for in the estimation procedure (for the single equation OLS estimates R_G^2 = .645).[17)] On the other hand, estimates of r do not seem to capture the main phenomenon predicted by the variance hypothesis. All in all, compared with single equation estimates the SUR estimates are more unfavourable to the variance hypothesis to the point of suggesting its rejection in the light of real output equation evidence.

Finally, consider the performance of the inflation rate equation (5). In this connection we neither used system methods of estimation nor made serial correlation correction. The results from using the non-linear least squares estimation with the restriction that the sum of foefficients of Δx_t and Δx_{t-1} is unity are reported in Table 8. They do not seem to conform to the variance hypothesis; e.g. Norway and Sweden have lower estimates of r than that of Iceland. Moreover, the goodness-of-fit statistics show no clear pattern.

Table 7. Two-state Aitken estimates of output equation

country	r* (ρ=0)	r*(Parks)	ρ
Denmark	.4541 (.1251)	.5824 (.1199)	.470
Finland	.3132 (.1018)	.3354 (.1127)	.434
Iceland	.2058 (.1086)	.3079 (.1211)	.321
Norway	.2617 (.0622)	.2696 (.0597)	-.048
Sweden	.0701 (.1209)	.0666 (.1346)	.191
	R_G^2=.712	R_G^2=.745	

Numbers in parentheses are (asymptotic) standard errors.

Table 8. Estimates of inflation rate equation

country	r*	$\bar{R}^2$	D-W	χ^2
Denmark	.7159 (5.5)	.655	1.90	.13
Finland	.4779 (3.8)	.691	1.98	.13
Iceland	.4291 (2.8)	.738	1.29	1.17
Norway	.1585 (0.5)	.826	1.93	.82
Sweden	.2242 (2.5)	.796	1.46	4.33

Critical t-values: $t_{.05,21} = 1.721$, $t_{.01,21} = 2.831$, lower and upper bounds of D-W statistics: $d_{L.05} = 1.12$, $d_{U.05} = 1.66$. With χ^2-statistics we test the hypothesis of no serial correlation of residuals, $\chi^2_{.10,3} = 6.25$.

Concludingly, the Scandinavian cross-country annual data do not seem to point to the supporting evidence for the variance hypothesis. The most favourable evidence on single equation estimates of output equation gives mixed results, and using Zellner's SUR estimation technique both makes the performance of output equations better and fails to support the variance hypothesis.18)

All estimates presented thus far have been based upon the assumption of unit-price elasticity of aggregate demand. The estimate of r is biased upward (downward) if the price elasticity of aggregate demand is greater (less) than unity (Arak [1]). The elasticities can be computed from the following equations

(4') $y_{c,t} = \beta_1 y_{c,t-1} + v_t$

(5') $\Delta P_t = \beta_2 + \beta_3 \Delta y_{c,t-1} + w_t$

where v_t and w_t are residuals. The price elasticity of aggregate demand is now $e_p = (1-\beta_1)/\beta_3$ (i.e., $y_t = -e_p P_t + x_t$), and its OLS estimates are reported in Table 9.

Table 9. OLS estimates of price elasticities

country	Denmark	Finland	Iceland	Norway	Sweden
$1-\beta^*_1$	.231 (1.6)	.504 (2.7)	.395 (2.4)	.136 (1.5)	.121 (0.8)
β^*_3	.358 (2.2)	.571 (1.9)	.016 (0.1)	.333 (1.6)	.164 (0.5)
e^*_p	.645	.883	24.688	.408	.738

Critical t-values: $t_{.05,23} = 1.714$, $t_{.05,24} = 1.711$.

In of case of Iceland, Norway and Sweden the price elasticity of aggregate demand connot be determined very accurately from (4') and (5'). But considering only the numerical values of e^*_p all of them - excluding that of Iceland - are below unity. From this point of view estimates of r are upward biased for Iceland, while downward biased for other countries. Accounting for this evidence would shift estimates of r towards the values predicted by the variance hypothesis, and might explain different results obtained by using periodic time-series data and cross-country data.[19)]

IV. CONCLUDING REMARKS

We have tested for the variance hypothesis on the output-inflation trade-off by using Finnish time-series quarterly data over the period 1948-1977 and cross-country annual data on Scandinavian countries over the period 1950-1976. The periodic time series data from Finland turned out to be in conformity with the variance hypothesis. Turning to Scandinavian cross-country data, some of the single equation estimates gave mixed results. The error terms across countries, however, displayed in quite many cases significant contemporaneous correlation so that system estimation techniques seemed appropriate. Experiments with Zellner's seemingly unrelated regression estimation procedure unambiguously failed to point to the supporting evidence for the variance hypothesis.

Finally, we might mention some agenda for further research. A problem with performed tests - and with those carried out by Lucas [18] - lies in the use of a very simple proxy for the trend component of supply, y_{nt}. Obviously, it would be worthwhile to experiment with more sophisticated proxies.

The notion behind the variance hypothesis that fluctuations in employment and output are due to errors in forecasting prices is inconsistent with the more usual notion of involuntary unemployment and has been subject to some criticism. It is well-known that the observed tradeoff between output and inflation can also be explained by the non-market-clearing paradigm of Barro-Grossman-Malinvaud type. Nowadays the current opinions on the relative usefulness of incomplete information and non-market-clearing paradigms to account for the output-inflation tradeoff are strikingly splitted (compare Modigliani [24] with Lucas [21]). Added emphasis should thus be given to the task of designing tests which would make it possible to distinguish between paradigms. This is reinforced by the fact that important policy questions are at stake.

FOOTNOTES:

1) By guessing the manner in which people form their expectations in practice and trying to find some quantitative representation of this behaviour, e.g. the adaptive expectations hypothesis promised unlimited real output gains from a well-chosen inflationary policy.

2) Moreover, since the behaviour of agents depends also on their perception of policy rules, the private sector behavioural relationships are not invariant to policy feedback rules (Lucas [17]).

3) McCallum [22] has demonstrated that price level stickiness by itself is not always sufficient to invalidate the neutrality result. But it is also easy to construct quite plausible models with sluggish price and wage adjustment that leave scope for real effects of policy feedback rules (see a survey by Buiter [6]). This need to distinguish carefully between REH and the models to which it is applied is strikingly illustrated also by Neary and Stiglitz [25] in showing how rational constraint expectations in Barro-Grossman-Malinvaud type disequilibrium model actually raise the value of government multiplier! For a further discussion of the role of market-clearing in RE models, see Arrow [2].

4) One such test have been recently constructed by Neftci and Sargent [26].

5) Vining and Elwertowski [35] have found a positive association between s^2 and q^2 which they consider as a contradiction to Lucas' model. Cukierman [9] has shown, however, that their findings are fully consistent with Lucas' framework; the variability in the general level of prices and in the relative prices may not be independent, but both may depend on the variability of nominal incomes (for some further developments and empirical evidence, see Cukierman and Wachtel [10], Parks [28]).

6) This particular characterization of 'shocks', x_t, is used partly for simplicity. Moreover, calculating the autocorrelation function for Finnish time-series data suggests that the assumption that Δx_t follows a random walk is not a very inaccurate one. For the Scandinavian cross-country data the number of observations does not make it reasonable to test for the behaviour of Δx_t.

7) Cukierman [8] has generalized the Lucas model to cover the situation in which agents are allowed to collect price information from several markets and to decide on the optimal amount of information collection. Not surprisingly, this modification is found to reinforce the variance hypothesis.

8) On the basis of Table 1 the existence of a stable Phillips curve may be directly suspected. If a stable Phillips curve would exist, then $var(y_{c,t})$ and $var(\Delta x_t)$ should move together, which does not seem to be the case. Moreover, there is no systematic relationships between average real growth ($\overline{\Delta y_t}$) and average inflation rate ($\overline{\Delta P_t}$).

9) This was used since the output equations are serially correlated in terms of Durbin's m-statistic (see Spencer [33]).

10) The Chow-test to check the null hypothesis that the periodic equations have identical coefficients were used. The null hypothesis was rejected for periods I and IV at the 1 per cent level, while it could not be rejected for periods II and III.

11) How do we know a priori that the countries have different policy regimes? They might have the same regime with only different realizations of the underlying stochastic process. Testing for the equality of sample variances of Δx_t for various countries indicated that Finland and Iceland on the one hand, and Denmark, Norway and Sweden on the other hand formed groups for which sample variances of Δx_t were significantly different.

12) Sweden was already a 'problem' in Lucas' original tests (Lucas [18], p. 332). In Lucas' data the Spearman rank correlation coefficient between var(dx_t) and ($1/r^*$) is .508, which is not significantly different from zero at the 1 per cent level. Dropping Sweden from data gives the rank correlation coefficient r_s = .631 which does differ significantly from zero at the 1 per cent level.

13) See Srivastava and Dwivedi [34] for a survey of estimating SUR-equations.

14) See Kmenta and Gilbert [15] for small sample properties of various ways to estimate SUR-equations with no serial correlation.

15) We also made computations by restricting r to be equal for four countries, while letting one r to be determined freely. Only in the case of Denmark - when its r-coefficient was allowed to differ from the common coefficient for other four countries - the performance of output equations turned out to be better in terms of χ^2-statistic at the 5 per cent level compared with the performance of output equations with an equal r-coefficient for all countries.

16) See Kmenta and Gilbert [14] for small sample properties to estimate SUR-equations when disturbances are both serially and contemporaneously correlated. The efficiency gain from system estimation did not seem to be very sensitive to the accuracy of the estimated serial correlation coefficients.

17) If disturbances are neither contemporaneously nor serially correlated then R_G^2 is a weighted average of the R^2 obtained from each equation estimated by OLS (see Dhrymes [11], p. 245).

18) Hamburger and Reisch [13] have estimated Lucas' model for some western European countries. They were not, however, interested in the variance hypothesis, but in comparing its performance with more traditional models which focus on the explanation of aggregate demand.

19) Lucas has defended the unit-price elasticity assumption as follows: "My strategy was based on the hope that the sample would exhibit enough cross-country variation to overcome what must necessarily

be large measurement errors on the supply elasticities for each country individually" (Lucas [20], p. 731). But if that is the case, then the explicit errors in variables specification should be used. We also estimated the price elasticities of aggregate demand for various subperiods with Finnish quarterly data, but could not reject the hypothesis that they are equal for subperiods. Finally, we should mention the following possibility: Although the variance hypothesis might very well be true, different countries might have different relative price parameters e in $r = we/(1+we)$ thereby blurring the evidence of cross-country estimation of r. But we have no constructive suggestions to deal with this problem.

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