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NEW INTERNATIONAL EVIDENCE ON OUTPUT-INFLATION TRADEOFFS - A NOTE***

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Abstract

This note presents some new international evidence for the variance hypothesis by Lucas according to which the output-inflation tradeoff varies inversely with the variance of nominal aggregate demand. The annual cross-country data on 18 countries are used. Over the period 1952-1967 the variance hypothesis gets support, and allowing for serial and contemporaneous correlation of error terms leaves the results practically unaffected. Over the period 1952-1977, however, the variance hypothesis gets stronger support, which is even reinforced by allowing for serial and contemporaneous correlation of error terms between the OLS residuals.

1. INTRODUCTION

When testing rational expectations hypothesis (REH) the basic difficulty lies in the fact that REH cannot generally be tested in isolation, but in the context of some other hypotheses (see e.g. McCallum [8])[1]. Indeed; Sargent [9] has shown that it is literally impossible to distinguish between Keynesian and classical macroeconomic structures using only parameter estimates from a single policy regime. One way to design proper tests is just to find periods across which the policy regimes differ and to test the invariance of alternative models across regimes[2]. Along these lines the use of cross-èquation restrictions to test 'natural' rate-rational expectations model has been suggested by Lucas [6]. According to the Lucas hypothesis the output-inflation tradeoff varies inversely with the variance of monetary and fiscal innovations. Lucas tested this hypothesis for 18 countries over the period 1952-1967 and found some support to it. This test has widely been considered to be the most important piece of evidence in favour of the 'new classical macroeconomics' (see e.g. Sargent [11], p. 330).

This note contains further tests of the Lucas hypothesis on output-inflation tradeoffs in the following respects: First, while using the same sample of countries as Lucas we test for the hypothesis over the longer period 1952-1977. In the light of high inflation rates in seventies compared with the earlier period it is interesting to look at the performance of the hypothesis under various circumstances. Second, Lucas estimated equations one-by-one by using OLS. It might be the case, however, that error terms across countries are contemporaneously correlated, in which case Zellner's seemingly unrelated system equation estimation technique (SUR) is the appropriate estimation procedure. Checking whether results are sensitive to the use of single-equation estimation technique is our second aim. Third, since a proxy for the trend component of supply used by Lucas may not be very statisfactory over long periods like 1952-1977, we introduce population as an additional explanatory variable. Finally, we report the single-equation estimation results also by accounting for the first-order serial correlation correction.

2. EMPIRICAL RESULTS

In Lucas's model the aggregate supply, y_t, is a product of a long run trend component, y_{nt}, and a cyclical component y_{ct} so that (in log terms) we have $y_t = y_{nt} + y_{ct}$. Under certain assumptions (see e.g. Sargent [11], pp. 325-330) the following aggregate supply function can be derived $y_{ct} = \theta(p_t - \bar{p}_t) + \lambda y_{c,t-1}$ where p_t denotes (actual) prices which are assumed to be normally distributed with mean $\bar{p}_t$, and where θ depends on the ratio of variances of relative to general price movements.

To complete the model Lucas postulates an aggregate demand function of the form $x_t = y_t + p_t$, where x_t = nominal aggregate demand, an exogenous shift variable. Assuming that Δx_t is a sequence of independent, normal variates with mean u_x and combining the supply function with x_t and with the assumption of rational price expectations gives the following equation for the equilibrium value of real output.

$$(1) \qquad y_{ct} = - \pi u_x + \pi \Delta x_t + \lambda y_{c,t-1}$$

where $\pi = \theta/(1+\theta)$. According to the variance hypothesis the coefficient π is inversely related to the variability of nominal aggregate demand. More specifically, π should be inversely related to var (Δx_t)[3].

Turn now to empirical results. The main test is to fit (1) and compare the estimates of π with the variances of Δx_t. Attention is also paid to the explanatory power of (1); its fit should be 'poor' if GNP has shown high variability and vice versa.

The results over the period 1952-1967 for the sample of 18 countries (used also by Lucas) are reported in Table 1. The OLS results are presented without and with the first-order serial correlation correction. In implementing the SUR-technique the maximum likelihood estimation method were used[4]. Similar estimations have been carried out over the period 1952-1977 and they are reported in Table 2[5].

The performance of the variance hypothesis on output-inflation tradeoff can be evaluated by means of Spearman rank correlation coefficients (r_s). These coefficients are computed with respect to estimates of π and to $(1/\text{var}(\Delta x_t))$ on the one hand, and with respect to R^2 and to $(1/\text{var}(\Delta x_t))$ on the other hand (see the last rows in Tables 1 and 2). For the period 1952-1967 the rank correlation coefficient between estimates of π and $(1/\text{var}(\Delta x_t))$ is just significant at the 5 per cent level for the OLS estimates, while it is slightly below the 5 per cent significance level both for the OLS estimates with the first-order serial correlation correction obtained by the Hildreth-Lu procedure and for the SUR estimation. The rank correlation coefficient between R^2 and $(1/\text{var}(\Delta x_t))$ follows a rather similar pattern. Thus allowing for serial and contemporaneous correlation of error terms leaves the results practically unaffected.

As far as results for the longer period 1952-1977 are concerned, the rank correlation coefficient between estimates of π and $(1/var(\Delta x_t))$ is significant at the 1 per cent level in all cases; allowing for serial and contemporaneous correlation of error terms makes r_s even higher[6). On the other hand, r_s between R^2 and $(1/(\Delta x_t))$ is practically unaffected by the first-order serial correlation correction[7). Thus extending the data sample to cover the years 1968-1977 reinforces the case for the variance hypothesis. Indeed, for the group of 18 countries the performance of the real output equations lies in striking conformity with the Lucas hypothesis over the period 1952-1977.

The previous estimates are based on the simple specification on the trend component of output y_{nt} = a+bt, which may not be a very satisfactory way of modelling y_{nt} for long periods like 1952-1977. Therefore, we introduced the (estimates of midyear) population P as an additional explanatory variable so that $y_{nt} = a+bt+cP_t$ (P is presented in log terms). The corresponding results without and with the first-order serial correlation correction are presented also in Table 2. They turn out to be even slightly better than those obtained by using y_{nt} = a+bt. GLS estimates of (1) have been recorded in appendix. On the whole, equation (1) performs rather well; coefficient estimates of Δx_t and $y_{c,t-1}$ are of expected signs and mostly highly significant and only few cases display serial correlation according to the Durbin's h-statistic.

We also restricted estimates of r to be equal across countries in the subsamples I-IV used in the SUR-estimation as a sort weak test of the variance hypothesis. With the exception of Scandinavian countries the equality hypothesis could be rejected for both periods for groups I-III and for the period 1952-1977 for group V[8).

Table 1. Estimation results for the period 1952-1967

Country		OLS		Hildreth-Lu		FIML
	$var(\Delta x_t)$	$\hat{\pi}$	R^2	$\hat{\pi}$	R^2	$\hat{\pi}$
Argentina	.01563	.014	.009	-.007	.019	.017
Guatemala	.00067	.418	.319	.409	.347	.288
Honduras	.00117	.358	.366	.323	.384	.290
Paraguay	.07195	.007	.190	.059	.171	.024
Venezuela	.00100	.459	.695	.460	.695	.536
Canada	.00141	.752	.939	.763	.959	.780
Puerto Rico	.00065	.698	.943	.656	.949	.590
United States	.00066	.931	.955	.907	.978	.833
Austria	.00121	.308	.495	.275	.616	.350
Belgium	.00071	.502	.861	.517	.903	.611
West-Germany	.00075	.783	.893	.679	.895	.724
Italy	.00070	.593	.820	.563	.846	.377
Netherlands	.00105	.534	.730	.551	.756	.515
Denmark	.00082	.553	.749	.593	.758	.514
Norway	.00100	.578	.842	.613	.859	.581
Sweden	.00047	.275	.488	.291	.562	.450
Ireland	.00108	.412	.813	.341	.820	.722
United Kingdom	.00017	.509	.358	.459	.374	.707
r_s:		.402	.406	.391	.249	.372

Critical values for r_s: $r_s(.05,18) = .399$, $r_s(.01,18) = .564$

Data source: Yearbook of National Account Statistics

Table 2. Estimation results for the period 1952-1977

		OLS		Hildreth-Lu		FIML	OLS		Hildreth-Lu	
Country	var(Δx_t)	$\hat{\pi}_{(1)}$	$R^2_{(1)}$	$\hat{\pi}_{(1)}$	$R^2_{(1)}$	$\hat{\pi}_{(1)}$	$\hat{\pi}_{(2)}$	$R^2_{(2)}$	$\hat{\pi}_{(2)}$	$R^2_{(2)}$
Argentina	.13275	-.022	.194	-.045	.305	-.042	-.028	.142	-.040	.186
Guatemala	.00383	.211	.593	.213	.595	.375	.216	.562	.219	.563
Honduras	.00165	.268	.442	.316	.528	.461	.226	.248	.282	.323
Paraguay	.04751	.029	.598	.053	.730	.058	.044	.469	.081	.493
Venezuela	.00883	.033	.771	.044	.778	.050	.053	.696	.060	.720
Canada	.00172	.464	.751	.606	.826	.613	.341	.449	.597	.701
Puerto Rico	.00059	.781	.877	1.033	.900	.822	.818	.885	.941	.894
United States	.00073	.612	.715	.934	.894	.766	.627	.727	.935	.893
Austria	.00107	.346	.570	.352	.655	.410	.356	.543	.358	.649
Belgium	.00131	.337	.776	.627	.842	.549	.298	.687	.634	.811
West-Germany	.00068	.774	.889	.725	.928	.558	.735	.802	.734	.821
Italy	.00198	.028	.672	.578	.864	.260	.009	.629	.481	.790
Netherlands	.00094	.456	.663	.580	.795	.511	.462	.670	.572	.775
Denmark	.00086	.459	.623	.749	.762	.547	.488	.573	.661	.632
Norway	.00127	.333	.814	.328	.815	.408	.209	.512	.298	.582
Sweden	.00066	-.062	.645	.355	.751	.439	.144	.401	.441	.582
Ireland	.00321	.202	.875	.183	.894	.201	.042	.638	.080	.759
United Kingdom	.00184	-.069	.176	-.045	.183	-.023	.026	.022	.032	.031
r_s:		.629	.404	.802	.402	.798	.785	.389	.821	.572

(1) refers to $y_{nt} = a + bt$, and (2) to $y_{nt} = a + bt + cP_t$. Critical values for r_s: $r_s(.05,18) = .399$, $r_s(.01.18) = .564$
Data sources: Yearbook of National Account Statistics and UN Demographic Yearbook.

Finally, values of π and var(Δx_t) (see Table 2) across countries, 1952-1977, are described in Figure 1. Their negative relationship is clearly to be seen even when the high var(Δx_t) -countries, Argentina, Paraguay and Venezuela, are dropped. Thus it is incorrect to argue that Latin American countries, which have large variations in nominal income and a low correlation between output and changes in nominal income, dominate the sample in the sense that the variance hypothesis would break down without them.

This evidence is reinforced by an informal comparison between the variances of Δp_t and y_{ct} (see Figure 2). The stable output-inflation tradeoff would imply their positive relationship. Such a clear pattern cannot, however, be discerned[9)].

Figure 1: Values of π and $var(\Delta x_t)$ across countries 1952-1977

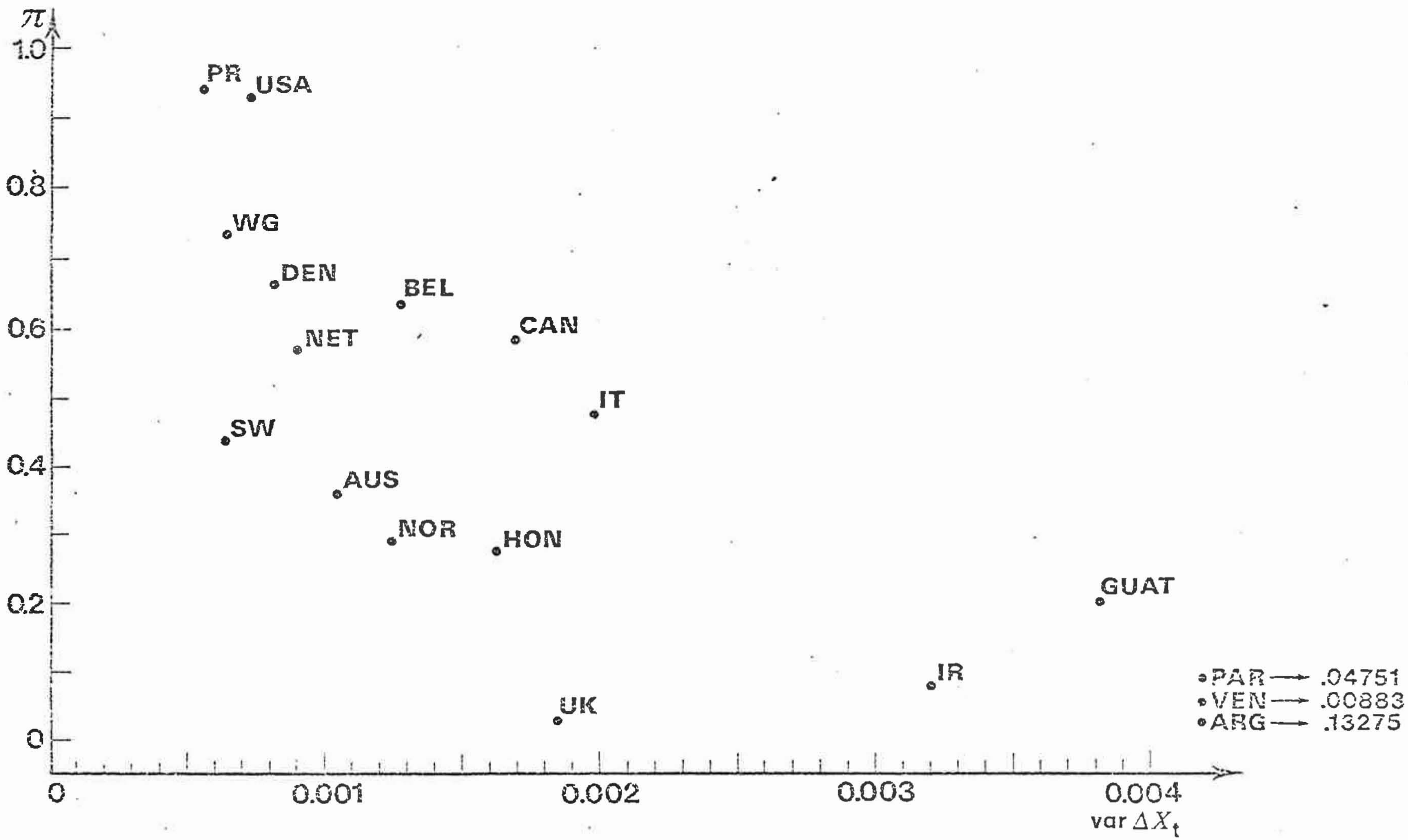
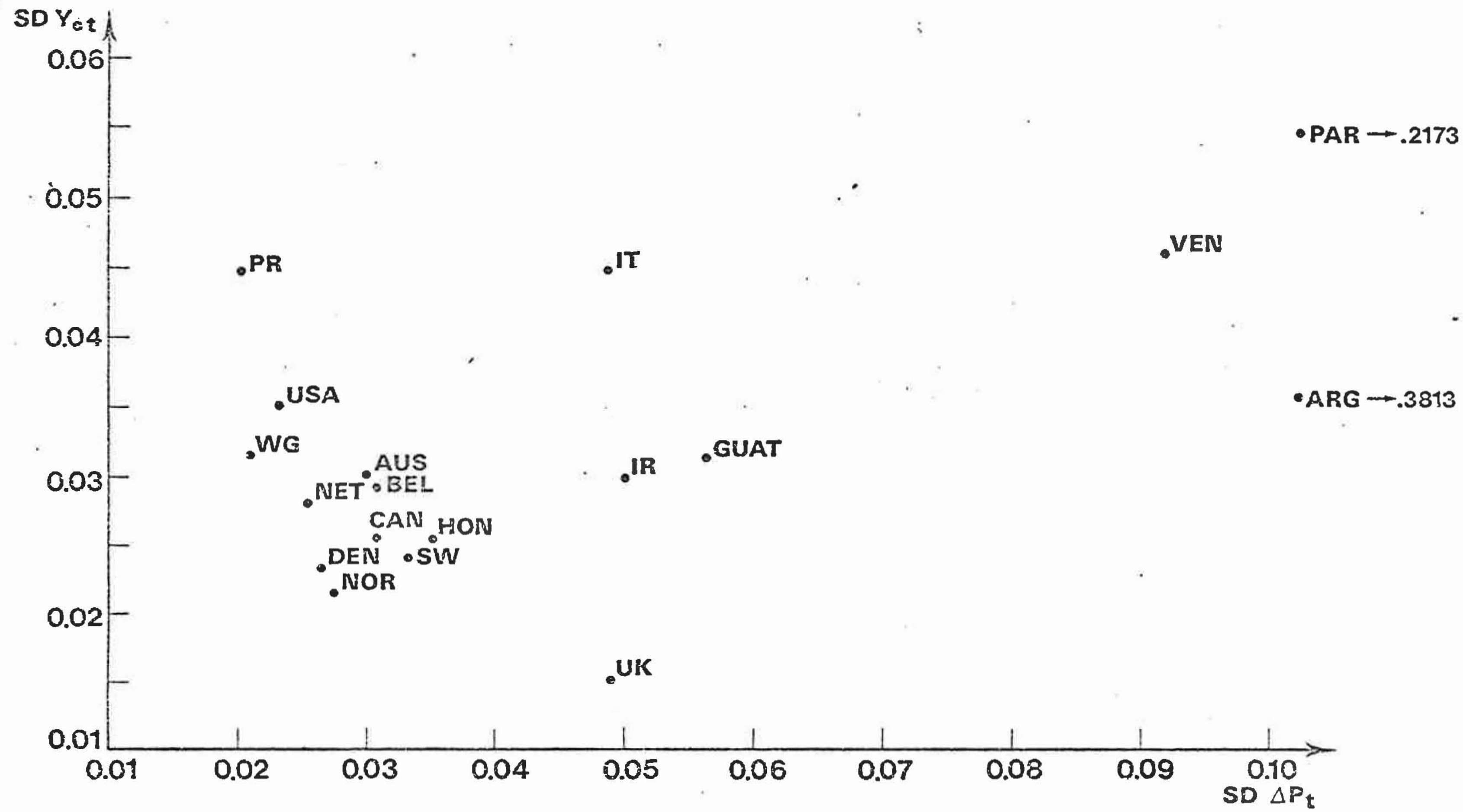


Figure 2. Standard deviations of Δp_t and y_{ct} across countries 1952-1977



Appendix: GLS estimates of (1) with the population proxy, 1952-1977

Country	Constant	Δx_t	$y_{c,t-1}$	R^2	h	ρ
Argentina (ARG)	.0105 (0.69)	-.0397 (1.64)	.0674 (0.35)	.1862	2.5977	.454
Austria (AUS)	-.0548 (2.00)	.3580 (4.55)	.2405 (1.69)	.6494	.4173	.649
Belgium (BEL)	-.0515 (2.73)	.6344 (7.13)	.3458 (2.92)	.8111	2.6260	.900
Canada (CAN)	-.0550 (3.30)	.5973 (7.11)	.4787 (3.55)	.7013	.9224	.843
Denmark (DEN)	-.0629 (4.95)	.6608 (5.54)	.3900 (2.73)	.6324	1.2170	.632
West Germany (WG)	-.0592 (5.62)	.7343 (6.86)	.4712 (4.65)	.8207	.2436	.489
Guatemala (GUAT)	-.0183 (2.44)	.2191 (3.00)	.5508 (3.85)	.5629	1.2352	.091
Honduras (HON)	-.0189 (1.77)	.2822 (2.59)	.2607 (1.29)	.3226	..	.463
Ireland (IR)	-.0092 (0.83)	.0805 (1.12)	.6378 (3.90)	.7590	2.3241	.645
Italy (IT)	-.0729 (2.31)	.4806 (4.45)	.5264 (2.83)	.7896	4.5869	.900
Netherlands (NET)	-.0565 (3.98)	.5718 (4.11)	.4915 (5.40)	.7745	.1519	.646
Norway (NOR)	-.0288 (2.54)	.2982 (3.11)	.3396 (1.76)	.5821	2.5144	.610
Paraguay (PAR	-.0096 (1.19)	.0814 (2.45)	.8052 (6.76)	.4933	1.2627	-.492
Puerto Rico (PR)	-.0874 (5.41)	.9412 (5.43)	.7761 (8.52)	.8935	.5154	.403
Sweden (SW)	-.0539 (1.91)	.4408 (3.29)	-.0034 (0.02)	.5817	..	.900
United Kingdom (UK)	-.0033 (0.43)	.0318 (0.41)	.0322 (0.15)	.0313	..	.141
United States (USA)	-.0646 (3.68)	.9347 (11.01)	.7061 (6.59)	.8928	11.6561	.900
Venezuela (VEN)	-.0007 (0.08)	.0603 (1.20)	.7070 (5.15)	.7196	.3584	.347

Critical t-values: $t_{.05,23} = 1.71$, $t_{.01,23} = 2.50$.

FOOTNOTES:

1) A notable exception is a recent study by Leiderman [5], in which REH is tested with and without a 'natural' rate hypothesis.

2) One such test have been recently constructed by Neftci and Sargent [9].

3) In Lucas [6] the equation for the equilibrium value of the inflation rate was also derived. The real output and inflation rate equations, however, do not consist of a two-equation model as claimed in Lucas [6], but they are transforms of each other (see Lucas [7]). Kormendi and McGuire [3], however, have recently shown how this relationship might as well obtain even if there were no true relationship between the reduced form coefficients and the variability of nominal aggregate demand, in which case relationship would be statistically spurious.

4) It turned out that the OLS residuals were both serially and contemporaneously correlated. Durbin's h-test indicated serial correlation in 6 cases over the period 1952-1967 and in 11 cases over the period 1952-1977 at the 5 per cent significance level. On the other hand, 28 (52) from the total number of 306 coefficients of correlation between the OLS residuals turned out to be different from zero over the period 1952-1967 (1952-1977) at the 5 per cent significance level. Finally, the likelihood ratio tests indicated that the SUR estimation technique increased the fit of the real output equations; the following likelihood ratio test values were obtained for the groups of countries used above in SUR estimation (the groups of countries are in the same order as in Tables 1 and 2 and the test values are over the period 1952-1977):

group of countries	χ^2	critical test values 5 %/1 %
I	37.255	18.307/23.209
II	12.920	7.815/11.345
III	86.727	18.307/23.209
IV	31.352	7.815/11.345
V	5.235	3.841/ 6.635

Thus the hypothesis that the covariance matrix of residuals is diagonal can be rejected at the 1 per cent significance level for the groups of countries I-IV and at the 5 per cent significance level for the group of countries V.

5) In two cases - Guatemala and U.K. - we could not reproduce Lucas' OLS estimates over the period 1952-1967. In the connection of the SUR estimation we experimented with numerous alternative cross-country subsamples without any clear difference in results. Some experiments turned out to be even slightly more favourable to the variance hypothesis than those presented in Tables 1 and 2. So e.g. if the 8 American

countries are divided into the following subsamples, I' = (Argentina, Paraguay, Venezuela), II' = (Canada, Guatemala, Honduras, Puerto Rico, United States), then $r_s = .849$!

6) A reason for this might be the following: Over the period 1952-1967 (1952-1977) 108 (172) from the total number of 306 F-test statistics - computed to test for differences in country variances - were significant at the 5 per cent level thus indicating that the differences between $var(\Delta x_t)$ have increased in the seventies.

7) In the SUR-estimation no correction for serial correlation was made due to the lack of computer program.

8) Hamburger and Reisch [2] have estimated Lucas' model for some western European countries over the period 1953-1973. They were not, however, interested in testing for the variance hypothesis, but in comparing the performance of Lucas' model with more traditional models which focus on the explanation of aggregate demand. Only the OLS estimates for the real output and inflation rate equations are reported in [2]. We have also tested for the variance hypothesis with the annual cross-country data on Denmark, Finland, Iceland, Norway and Sweden over the period 1950-1976. Results fail to point to the supporting evidence for the hypothesis, particularly when Zellner's seemingly unrelated regression estimation procedure is used (see Koskela and Virén [4]). Recently, Froyen and Waud [1] have presented some cross-country evidence on the variance hypothesis over the period 1957-1976 with data sample: United Kingdom, Canada, United States, Italy, Japan, Belgium, Switzerland, France, Netherlands and West Germany. As a whole their findings are in agreement with those presented in this note.

9) We also estimated the equation $\Delta p_t = b_0 + b_1 y_{ct} + u_t$, where $y_{ct} = y_t - \hat{a} - \hat{b}t - \hat{c}p_t$. In all cases the "Phillips curve" hypothesis about the relationship between Δp_t and y_{ct} could be rejected at all conventional levels of significance.

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