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OUTPUT AND PRICE STABILITY

An International Comparison

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The paper examines the behavior of output and prices for 10 different countries within the context of a rational expectations model with wage contracts. The model predicts that there is a policy tradeoff between output stability and price stability, and the estimated correlations and dynamic interactions are generally consistent with the model. For each of the countries the model is fit to the data using a maximum likelihood procedure which constrains the estimates to satisfy the rational expectations hypothesis. The estimated coefficients are compared across the countries in order to determine whether the structure of the model is systematically related to economic policy, and therefore whether it is reasonable to assume that the parameters of the model are policy-invariant. It is shown how the model can be used to calculate optimal tradeoffs when the policy-invariance assumption is satisfied.

1. Introduction

One of the implications of recent 'contracting' theories of macroeconomic fluctuations is that there is a policy tradeoff between output stability and price stability – even in a world of rational expectations – and that this tradeoff plays an important role in explaining business cycle persistence [see Taylor (1979, 1980), for example]. Put simply, very accommodative aggregate demand policies may be able to smooth out business cycle fluctuations, but they generally result in large swings in inflation rates. On the other hand, less accommodative policies designed to achieve greater price stability tend to generate larger business cycle fluctuations. However, other theories, in particular those of Barro (1976), Lucas (1973), and Sargent and Wallace (1975), suggest that there is no stable tradeoff and that the degree of policy reaction to the state of the economy does not matter for the behavior of output and employment when people form expectations rationally. The difference between these conflicting theories can be traced to theoretical

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issues concerning the formation of wages and prices,[1] but the choice between them will depend on empirical as well as theoretical considerations.

The purpose of this paper is to provide some empirical perspective on this tradeoff question by examining aggregate demand policy rules and the resulting output and price dynamics for 10 countries: Austria, Canada, Denmark, Germany, Italy, Netherlands, Norway, Sweden, the United States and the United Kingdom during the 1954–76 period.[2] An international comparison of this kind is particularly useful for studying the effect of aggregate demand policy on the economy, for it permits larger variations in the policy rule than are normally observed in any single country. This advantage of international data has been emphasized by Sargent (1976) and was the primary rationale for the international cross-section test of the information-based theory of the Phillips curve conducted by Lucas (1973). More recently, Black (1977), Gordon (1977), Sims (1979) and others have used cross-sections of countries to examine various aspects of aggregate demand policy. Although more work along these lines is required before firm conclusions can be reached, the preliminary finding of this study is that the international data displays important empirical regularities which are consistent with those business cycle theories in which a policy tradeoff has an essential role. This finding is supported by statistical tests which tend to reject the hypothesis that output is exogenous to nominal variables. However, some evidence of a relationship between the policy and the structural parameters suggests that the full reaction of people to the policy rule may not be adequately accounted for by the theory. If so, then the tradeoff might show some sensitivity to policy-choice [see Lucas (1976)].

A technical by-product of this analysis is a test of the general applicability of econometric techniques for estimation and policy formation in rational expectations models. For each of the ten countries a simple rational expectations macromodel is estimated using a maximum likelihood procedure which constrains the parameters of the model to satisfy the rational expectations restrictions. The macromodel involves multiperiod forecasts of future inflation rates and business conditions. These multiperiod forecasts require certain factorization techniques and more complex non-linear restrictions than arise in the model estimated in Taylor (1979) by an alternative systems estimation technique. Nevertheless, the approach seems to work well as a general procedure for macromodel estimation.

The paper proceeds as follows: section 2 presents a descriptive statistical

[1]See McCallum (1979) for a summary of these issues.

[2]The ten countries are the same OECD countries studied by Lucas (1973) with the exception of Ireland and Belgium. This study does not consider the Latin American countries also considered by Lucas, in the hope of reducing the variability of economic structure compared with economic policy. A useful extension would be to consider these countries and others such as Japan and France.

analysis of annual data on real output and prices in the ten countries during the 1954–76 period. The elements of the contemporaneous variance–covariance matrix of detrended output and prices, and alternative univariate and bivariate descriptive time series models of their joint stochastic dynamics are compared across the different countries. These estimated statistical properties provide a set of 'stylized facts' about the behavior of the series which are not constrained by a particular macroeconomic theory. Prominent among these properties are a negative correlation between detrended price and output, a positive correlation between the rate of inflation and detrended output, and a persistence of output deviations from trend which is much stronger when output is considered as a univariate process than when output and prices are considered jointly. This last property is formalized by examining test statistics for the null hypothesis that output is exogenous to prices in each country.

Section 3 then presents a rudimentary stochastic macroeconomic model which displays many of the statistical properties observed in the series. Wage contracting is the primary feature of the model, and in this respect it differs from the rational expectations models of Barro, Lucas, and Sargent–Wallace which are based on flexible prices and information lags. The property of the model which we stress here is the policy tradeoff between output and price fluctuations.

In section 4 the model is estimated and the relationship between the structural and policy parameters across the countries is examined. This empirical comparison points to some of the reasons why different countries experience different types of business cycles and inflation instabilities. Output and price behavior can differ across countries either because economic policies are different or because the structures of the economies are different. An accommodative aggregate demand policy, for example, provides an explanation for relatively large price fluctuations and small output fluctuations, while more restrictive policies will alter the relative size. Divergences from this pattern may then be attributed to differences in economic structure.

2. Statistical properties of the data

This section provides estimates of the stochastic properties of output and prices in the ten countries from 1954–76. The stochastic system we focus on is the bivariate relationship between the logarithm of real GNP (or GDP) and the logarithm of the GNP (or GDP) deflator after a linear trend estimated over the sample period has been removed.[3] This particular method of detrending is, of course, arbitrary. A popular alternative method is to

[3]Output is measured by GNP in Austria, Canada, Germany, the Netherlands, and the U.S., and by GDP in other countries.

detrend by taking differences of both series, as suggested by Box and Jenkins (1971). Many economic studies take a mixed approach, detrending output using a deterministic trend while detrending prices by taking first differences.[4] Sims (1979) uses the deterministic detrending method for both series.

Let y_t be the detrended log of real output and p_t be the detrended log of the price level. Since these are logarithms they represent *percentage* fluctuations about trend and in this sense are unit free. The following descriptive statistical analysis focuses on the simple contemporaneous variance–covariance matrix and the moving average representation of the vector (y_t, p_t).

Table 1

Comparison of output–price variability and correlations, 1954–76.[a]

	σ_p	σ_y	$\rho(p_t, y_t)$	$\rho(p_{t+1}-p_t, y_t)$	$\rho(p_t-p_{t-1}, y_t)$
Austria	4.90	2.49	−0.293	0.358	0.123
Canada	8.77	2.47	0.131	0.408	0.265
Denmark	7.26	3.46	−0.690	0.303	−0.047
Germany	3.62	3.62	−0.619	0.439	0.097
Italy	10.86	3.66	−0.893	−0.099	−0.417
Netherlands	7.33	2.52	−0.128	0.392	0.123
Norway	7.52	1.71	0.293	0.284	0.413
Sweden	5.67	2.49	−0.783	0.014	−0.345
United Kingdom	11.84	1.92	−0.733	0.107	−0.442
United States	6.43	3.03	−0.618	0.303	−0.093

[a]*Note*: The variables y and p are deviations of the logarithms of real output and the output deflator from a linear trend. Output is measured by real GNP in Austria, Canada, Germany, the Netherlands, and the U.S., and by GDP in the other countries. The column headed by σ_p and σ_y are standard deviations (measured in percent) and the remaining columns are estimated correlation coefficients between the stated variables. Annual data is used throughout and is obtained from the International Financial Statistics (IFS) tape of April 1978.

2.1. Variances and contemporaneous correlations

The first two columns of table 1 compare the standard deviations of p_t and y_t for the ten countries. As these standard deviations will serve as our primary measures of price and output stability their variation across the countries is particularly important. What is most noticeable about these standard deviation pairs, is that prices show much more variation over the countries than output. Hence, if there is a negative tradeoff between σ_p and σ_y it would appear very flat (with σ_y on the vertical axis and σ_p on the horizontal) as judged by this scatter alone. It is true that Germany has the

[4]To the extent that unemployment is closely related to the deviations of employment from trend, Phillips (1958) detrended wages and employment in this way. Most large scale econometric models consider the first difference of the log of average hourly earnings as the basic variable. An exception is Dhrymes (1978) where level of the wage rate is the basic variable. Limited experimentation with alternative detrending techniques such as first differences and linear splines did not point to any serious deficiencies with the simple linear trend. But because of the near non-stationarity of the detrended price, further robustness studies would be useful.

most stable price level and a relatively high variability of output and that the U.K. has the least stable price level and a relatively low variability of output. But if this data is to give any evidence for a tradeoff, there must be some explanation for countries such as Italy – either inefficient aggregate demand policy or unfavorable economic structures.

The third column of table 1 gives the correlation coefficient between p_t and y_t. For all countries except Canada and Norway this correlation is negative and suggests that supply shocks have been dominating during the sample period. (The correlation is also negative, though smaller in absolute value, for all countries except the Netherlands from 1954–68.) The last two columns of table 1 list the correlation between the inflation rate and detrended output. If this is positive, then Phillips curve effects would be evident in the data. When the inflation rate is measured over the year following the output fluctuation, there is a positive correlation (with the exception of Italy) despite the negative correlation between the levels. However, the more contemporaneous correlation is smaller in all countries and frequently negative.

2.2. *Stochastic dynamics*

A simple way to summarize the persistence of business cycles in these countries is through univariate moving average models for detrended output. In table 2, second-order autoregressions are presented for each country along with the mean distance between peaks as implied by the autoregression. For each country the second-order term enters with a negative sign. This creates a lagged effect of shocks on output which either rises for the first few periods – as in Italy and Sweden – or diminishes more slowly in the first few periods than would be implied by simple geometric decay.[5] This is shown in the implied moving average representation presented in table 3 for the lags up to eight years. The persistence of output deviations from trend is shortest for Austria, Netherlands and the U.K., and longest for Sweden and Italy. This also corresponds to the ranking of persistence in terms of mean distances between peaks.

A more complete way to summarize the stochastic dynamics of these two series is through the bivariate moving average representations.[6] These were obtained for each country in the sample by first estimating the bivariate

[5]Quarterly data for the U.S. indicates that the lagged effects rise for the first few periods before declining. Annual data has a smoothing effect which cuts through this hump, but still gives a substantially more delayed decay than is implied by a first order autoregression.

[6]Sargent (1978) and Sims (1979) have used the multivariate moving average representation to characterize macroeconomic data in two recent studies. We do not orthogonalize the shocks because we are interested in interpreting the moving average representation using a simple structural macroeconomic model where the shocks need not be orthogonalized. Nevertheless, the correlation between shocks and the relative magnitude of the standard deviation of these shocks (given in table 6) should be taken into account when interpreting the results and working through stochastic simulations.

Table 2

Estimates of second-order autoregressive models for real output, 1956–76.[a]

$$y_t = a_1 y_{t-1} + a_2 y_{t-2} + e_t$$

	a_1	a_2	SSR	Mean distance between peaks $2\pi/\cos^{-1}(-0.5(1-a_1+a_2))$
Austria	0.711	−0.387	0.00601	4.13
Canada	0.983	−0.459	0.00508	4.66
Denmark	0.960	−0.271	0.01144	4.32
Germany	0.958	−0.402	0.00964	4.52
Italy	1.109	−0.382	0.01056	4.75
Netherlands	0.740	−0.525	0.00706	4.37
Norway	0.853	−0.221	0.00197	4.10
Sweden	1.270	−0.463	0.00400	5.25
United Kingdom	0.741	−0.463	0.00515	4.28
United States	0.930	−0.379	0.00922	4.44

[a]*Note*: The variable y_t is equal to the deviation of the log of output (real GNP or GDP) from trend. The mean distance between peaks is given in years and reflects the stochastic features of the second order difference equations. See Yule and Kendall (1965, p. 657).

Table 3

Moving average representation for output.

$$y_t = \sum_{i=0}^{\infty} \psi_i \varepsilon_{t-i}$$

	Austria	Canada	Denmark	Germany	Italy
ψ_0	1.00	1.00	1.00	1.00	1.00
ψ_1	0.71	0.98	0.96	0.96	1.11
ψ_2	0.12	0.51	0.65	0.52	0.85
ψ_3	−0.19	0.05	0.37	0.11	0.52
ψ_4	−0.18	−0.19	0.17	−0.10	0.25
ψ_5	0.06	−0.21	0.07	−0.14	0.08
ψ_6	0.03	−0.12	0.02	−0.10	−0.01
ψ_7	0.04	−0.02	−0.00	−0.03	−0.04
ψ_8	0.02	0.03	−0.01	0.01	−0.04

	Netherlands	Norway	Sweden	U.K.	U.S.
ψ_0	1.00	1.00	1.00	1.00	1.00
ψ_1	0.74	0.85	1.27	0.74	0.93
ψ_2	0.02	0.51	1.15	0.09	0.49
ψ_3	−0.37	0.24	0.87	−0.28	0.10
ψ_4	−0.29	0.10	0.58	−0.25	−0.09
ψ_5	−0.02	0.03	0.33	−0.05	−0.12
ψ_6	0.14	0.00	0.15	0.07	−0.08
ψ_7	0.11	0.00	0.04	0.08	−0.03
ψ_8	0.01	0.00	−0.02	0.03	0.00

Table 4

Estimates of bivariate autoregressive models for output and prices, 1956–76.

$$\begin{pmatrix} y_t \\ p_t \end{pmatrix} = \begin{pmatrix} a_{11} a_{12} \\ a_{21} a_{22} \end{pmatrix} \begin{pmatrix} y_{t-1} \\ p_{t-1} \end{pmatrix} + \begin{pmatrix} b_{11} b_{12} \\ b_{21} b_{22} \end{pmatrix} \begin{pmatrix} y_{t-2} \\ p_{t-2} \end{pmatrix} + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix}$$

	A		B	
Austria	0.688	−0.165	−0.308	0.079
	0.291	1.533	0.005	−0.659
Canada	0.984	−0.073	−0.402	0.061
	0.420	1.912	−0.410	−0.966
Denmark	0.260	−0.540	0.097	0.197
	0.350	1.778	0.085	−0.754
Germany	0.716	−0.359	−0.335	0.036
	0.290	1.447	−0.212	−0.570
Italy	0.333	−0.515	−0.037	0.314
	−0.159	2.055	−0.859	−1.469
Netherlands	0.740	−0.004	−0.459	−0.069
	0.222	1.484	0.252	−0.598
Norway	0.681	0.127	−0.138	−0.167
	0.322	1.447	0.164	−0.559
Sweden	0.689	−0.441	−0.015	0.290
	0.736	2.041	−0.891	−1.155
United Kingdom	0.259	−0.126	−0.503	−0.015
	1.222	1.877	0.402	−0.810
United States	0.575	−0.310	−0.220	0.079
	0.266	1.842	−0.26	−0.852

second-order autoregression given by

$$z_t = Az_{t-1} + Bz_{t-2} + u_t, \tag{1}$$

where A and B are 2×2 matrices, $z_t \doteq (y_t, p_t)'$, and u_t is a serially uncorrelated stochastic vector with zero mean. Estimates of A and B are given in table 4 and the implied infinite moving average representation.

$$z_t = \sum_{i=0}^{\infty} \Psi_i u_{t-i}, \tag{2}$$

with $\Psi_0 = I$ is given in table 6.

Two important characteristics of the data are evident in table 6. First, the own-persistence of output (that is, the lagged effects of shocks to the output equation) is much shorter than was indicated by the univariate representation in table 3 – exceptions are Austria, Canada, and the Netherlands.

Since the univariate model mixes output and price shocks together, this reduction in the level of persistence is one measure of the amount of business cycle persistence which was caused by policy attempts to stabilize inflation. Second, the impact of price shocks on output (ψ_{12}) is negative, while the impact of output shocks on prices (ψ_{21}) is positive. The stochastic behavior of output and prices are interrelated in this particular way for almost all countries in the sample.

Table 5

F-tests for exogeneity of output and prices, 1956–76.[a]

	$H_0: a_{12}=b_{12}=0$ (output exogenous)	$H_0: a_{21}=b_{21}=0$ (price level exogenous)
Austria	0.55	1.03
Canada	0.14	2.25
Denmark	5.55	0.98
Germany	2.16	1.29
Italy	11.77	5.66
Netherlands	0.49	1.47
Norway	2.28	0.46
Sweden	5.05	2.53
United Kingdom	5.00	2.86
United States	3.07	2.24

[a]*Note*: Significance points are 3.68, 4.77, and 6.36 for $F(2,15)$ at the 5, 2.5, and 1 percent levels, respectively. The results are very similar if 2 times each value is treated as a χ^2 variable with 2 degrees of freedom.

A more formal way of examining this dynamic interrelation between prices and output is through tests of the null hypotheses (1) that output does not appear in the bivariate autoregression for prices, and (2) that prices do not appear in a bivariate autoregression for output. Hypothesis (1) implies that the nominal variable in this bivariate relation is exogenous to real variables, while hypothesis (2) implies that real variables are exogenous. The second is a characterization of the strong neutrality hypothesis against the non-neutrality hypothesis, if one makes the additional assumption that the real shocks are not serially correlated for other reasons [see Sims (1979)].

Such tests are shown in table 5. Although the power of these tests is likely to be weak with the limited sample size, there is very little evidence that the dynamic cross effect of output on prices is significant. With the exception of Italy one could not reject the null hypothesis that output deviations do not influence prices. (As will be shown below, however, this is not evidence that aggregate demand policy does not affect prices – such influences would be captured in the reduced form coefficients of the price equation.)

Table 6

Bivariate moving average representations for output and prices, 1956–76.[a]

Lag	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}
	Austria				*Canada*			
0	1.00	0.0	0.0	1.00	1.00	0.0	0.0	1.00
1	0.69	−0.17	0.29	1.53	0.98	−0.07	0.42	1.91
2	0.12	−0.29	0.65	1.64	0.54	−0.15	0.81	2.66
3	−0.22	−0.30	0.84	1.43	0.10	−0.20	0.96	3.20
4	−0.27	−0.22	0.80	1.02	−0.14	−0.21	0.87	3.54
5	−0.19	−0.12	0.60	0.55	−0.18	−0.19	0.64	3.66
6	−0.08	−0.02	0.33	0.44	−0.12	−0.15	0.37	3.59
7	0.05	0.04	0.09	−0.15	−0.03	−0.12	0.11	3.34
8	0.03	0.07	−0.08	−0.32	0.03	−0.07	−0.11	2.93
	($\sigma_1=1.63$,	$\sigma_2=1.55$,	$\rho=0.07$)		($\sigma_1=1.54$,	$\sigma_2=1.57$,	$\rho=0.41$)	
	Denmark				*Germany*			
0	1.00	0.00	0.00	1.00	1.00	0.00	0.00	1.00
1	0.26	−0.54	0.35	1.77	0.72	−0.36	0.29	1.45
2	0.02	−0.90	0.63	2.22	−0.07	−0.74	0.42	1.42
3	−0.25	−1.14	0.82	2.33	−0.33	−0.87	0.30	1.09
4	−0.39	−1.21	0.90	2.14	−0.35	−0.71	0.09	0.67
5	−0.45	−1.12	0.87	1.73	−0.16	−0.42	−0.07	0.32
6	−0.45	−0.92	0.74	1.16	0.03	−0.15	−0.13	0.12
7	−0.39	−0.64	0.54	0.54	0.12	−0.03	−0.10	0.03
8	−0.29	−0.32	0.30	−0.07	0.11	0.04	−0.05	0.01
	($\sigma_1=1.77$,	$\sigma_2=1.70$,	$\rho=-0.37$)		($\sigma_1=2.06$,	$\sigma_2=1.89$,	$\rho=-0.08$)	
	Italy				*Netherlands*			
0	1.00	0.00	0.00	1.00	1.00	0.00	0.00	1.00
1	0.33	−0.52	−0.16	2.06	0.70	−0.00	0.22	1.48
2	0.16	−0.92	−1.24	2.84	0.04	−0.08	0.74	1.60
3	0.63	−1.10	−2.62	3.40	−0.32	−0.16	1.15	1.48
4	1.17	−1.20	−3.31	3.79	−0.29	−0.19	1.20	1.18
5	1.50	−1.24	−4.70	3.94	−0.15	−0.17	0.95	0.78
6	1.68	−1.21	−5.31	3.75	−0.05	−0.11	0.59	0.37
7	1.76	−1.05	−5.57	3.18	−0.04	−0.06	0.25	0.01
8	1.73	−0.77	−5.38	2.24	−0.04	−0.01	0.34	−0.24
	($\sigma_1=1.40$,	$\sigma_2=1.60$,	$\rho=0.25$)		($\sigma_1=1.78$,	$\sigma_2=1.81$,	$\rho=0.12$)	
	Norway				*Sweden*			
0	1.00	0.00	0.00	1.00	1.00	0.00	0.00	1.00
1	0.68	0.13	0.32	1.45	0.69	−0.44	0.74	2.04
2	0.37	0.10	0.85	1.57	0.10	−0.91	1.12	2.69
3	0.21	0.01	1.27	1.52	−0.25	−1.20	0.89	2.85
4	0.11	−0.08	1.50	1.34	−0.24	−1.26	0.26	2.64
5	0.03	−0.14	1.53	1.07	−0.01	−1.14	−0.46	2.25
6	−0.05	−0.17	1.40	0.74	0.28	−0.95	−1.03	1.82
7	−0.12	−0.18	1.16	0.39	0.52	−0.74	−0.74	1.44
8	−0.16	−0.17	0.84	0.07	0.64	−0.57	−0.57	1.13
	($\sigma_1=0.85$,	$\sigma_2=2.65$,	$\rho=0.04$)		($\sigma_1=1.07$,	$\sigma_2=1.68$,	$\rho=-0.49$)	

Table 6 (continued)

Lag	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}
	United Kingdom				*United States*			
0	1.00	0.00	0.00	1.00	1.00	0.00	0.00	1.00
1	0.26	−0.13	1.22	1.88	0.58	−0.31	0.27	1.84
2	−0.59	−0.28	3.01	2.56	0.03	−0.67	0.62	2.46
3	−0.68	−0.36	4.05	2.89	−0.28	−0.93	0.90	2.79
4	−0.43	−0.35	4.10	2.80	−0.40	−1.06	1.06	2.80
5	−0.35	−0.30	3.61	2.34	−0.43	−1.05	1.09	2.54
6	−0.39	−0.24	2.85	1.61	−0.41	−0.94	1.00	2.03
7	−0.34	−0.14	1.83	0.71	−0.37	−0.74	0.81	1.36
8	−0.17	−0.03	0.55	−0.24	−0.29	−0.48	0.56	0.60
	($\sigma_1=1.21$, $\sigma_2=3.04$, $\rho=-0.43$)				($\sigma_1=1.77$, $\sigma_2=1.11$, $\rho=-0.15$			

[a]*Note*: The ψ_{ij} weights are the elements of the 2×2 matrix Ψ weights in the representation $z_t=\sum_{s=0}^{\infty}\Psi_s u_{t-s}$, where $z_t=(y_t,p_t)'$ and $u_t=(u_{1t},u_{2t})'$. Hence, the ψ_{11} weights represent the 'own-persistence' effects of output, ψ_{12} represents the effect of price shocks on output, ψ_{21} represents the effect of output shocks on prices, and ψ_{22} is the own-persistence effect of prices. The parameters σ_1 and σ_2 are the standard deviations of output and price shocks, respectively (in percent). The parameter ρ is the correlation coefficient between output and price shocks.

On the other hand, the effect of prices on output is stronger for most of the countries, and though insignificant in several instances, is very strong for Denmark, Italy, Sweden, and the United Kingdom. Given these results it is difficult to accept the hypothesis that no policy tradeoff exists for these countries. Moreover, as is shown in the following section, the apparent exogeneity observed in several countries may be the result of very accommodative aggregate demand policies which tend to insulate real variables from price shocks. In particular, Austria and the Netherlands, where output exogeneity is very difficult to reject, have very accommodative policies.

3. Structural model for output and prices

In this section we present a simple macroeconomic model which displays most of the statistical properties evident in the above analysis. The model is discussed in more detail in Taylor (1980), but is summarized here. An important feature of the model is the existence of overlapping nominal wage contracts. The staggering of these contracts generates a stochastic dynamic relationship between output and prices with persistence effects which may be quite long. The model exhibits a policy tradeoff and, as long as policy is not completely accommodative, implies that output should not be exogenous to price shocks. Since the annual data we examine in this paper suggest a period model of one year, the shortest length of contract we can consider is two years. To provide some uniformity to the intercountry analysis we will focus entirely on two year contracts in this paper. An investigation with quarterly data would be an extremely useful extension of this work for it would enable one to consider shorter contracts and to incorporate the institutional arrangements in each country. The two year assumption can only serve as an approximation.

As in section 2, let y_t and p_t be the deviation of the log of real output and prices from trend. Let x_t be the deviation from trend of the logarithm of the *contract* wage set in period t by firms who are negotiating with workers in period t. In keeping with the period model approximation we assume that the actual distribution of wage settlements can be approximated by one in which half the workers sign their contracts each year and in which the contracts last for two years. The basic model we postulate can be represented as

$$x_t = 0.5(x_{t-1} + \hat{x}_{t+1}) + 0.5\gamma(\hat{y}_t + \hat{y}_{t+1}) + \varepsilon_t, \tag{3}$$

$$p_t = 0.5\,(x_t + x_{t-1}), \tag{4}$$

$$y_t = -\beta p_t + v_t, \tag{5}$$

$$\varepsilon_t = u_t + \delta u_{t-1}, \tag{6}$$

$$v_t = \eta_t + \theta \eta_{t-1}, \tag{7}$$

where a hat over a variable represents its conditional expectation given information through period $t-1$, and where u_t and η_t are serially uncorrelated random shocks with zero means. Note that this implies that the random shocks ε_t and v_t are serially correlated according to a first-order moving average process.

Eq. (3) states that the contract wage set in period t is equal to the average wage expected to prevail during the contract period, plus a measure of expected excess demand during the contract period. If we assume that the excess demand for labor is proportional to the deviations of the log of output from trend, then y_t can serve as this excess demand measure expected at the end of period $t-1$. The parameter γ is a measure of the sensitivity of wages to excess demand, and ε_t is a serially correlated shock. Eq. (4) simply assumes that the aggregate price level is a proportional markup over the average wage in effect in a given period. The average wage is defined as the logarithm of the geometric average of the outstanding contract wages x_t and x_{t-1}. Since the variables are measured as deviations from trends and since we treat x_t as an unobservable, this markup assumes that all productivity changes are secular and are captured by the trend variable.

Eq. (5) can be derived directly from a quantity theory of aggregate demand $y_t + p_t = m_t + v_t$, with a monetary policy rule $m_t = (1-\beta)p_t$ for the deviations of the log of the money supply from trend. The result is the aggregate demand policy rule described by (5). It states that the policy response to a one percent deviation of the price level from its trend growth path is to let nominal aggregate demand increase by $1-\beta$ percent, that is, $(y_t+p_t=(1-\beta)p_t+v_t)$. If $\beta=0$ then aggregate demand policy is completely accommodative; when $\beta=1$ aggregate demand policy is monetarist with nominal income growth unresponsive to price changes. Note that the exclusion of output y_t from the monetary policy rule involves no loss of generality. If y_t appeared in the rule it could be substituted out via the quantity equation to give a relationship like (5).

This policy rule is an abstraction of a very complicated policy making process which includes the central bank, the government and the legislative body. However, it is similar to the characterization of policy which is frequently made in practical discussions. For example, in a set of recommendations to the Federal Reserve Board in 1977, Okun argued:[7]

> In principle, favorable or unfavorable surprises in inflation call for a reassessment of the target path for nominal GNP. And, in principle (technically, as long as both price stability and output are 'normal goods' in the social welfare function), the result will be a compromise. Bad news on inflation should lead toward less real growth, but not point for point; and hence it calls for *partial* accommodation in the growth of nominal GNP.

In the algebra of eq. (5), a 'point for point' reduction in y_t in response to an increase in p_t would correspond with $\beta=1$. 'Partial accommodation' is characterized by $0<\beta<1$.

An earlier example is Keynes' recommendation:[8]

> We might be able to moderate very greatly the amplitude of the fluctuations if it was understood that the time to deflate the supply of cash is when real balances are falling, i.e., when prices are rising out of the proportion to the increase, if any, in the volume of cash, and that the time to inflate the supply of cash is when real balances are rising, and not, as seems to be our present practice, the other way around.

This rule might be written $m_t=\gamma(m_t-p_t)$ with $0<\gamma<1$. Hence $\beta=(1-\gamma)^{-1}>1$. Accordingly Keynes' recommendation was for a more restrictive policy than 'point for point' monetarism.

The aggregate demand policy rule (15) involves a simultaneous reaction of policy to economic events during the current period. It may appear that we are going beyond the usual definition of feedback policy which states that policymakers respond to lagged variables. Hence, in order to facilitate comparison with other studies which consider only feedback policy, it is

[7]See Okun (1977).
[8]See Keynes (1924, p. 205).

useful to restrict the analysis to feedback policy. For the remaining analysis we replace p_t by $\hat{p}_t$ in eq. (5). Since $\hat{p}_t$ is predetermined, the aggregate demand policy rule is then of the feedback variety.[9]

In order to compare this model with the data it is necessary to obtain the reduced form solution of the structural model (3) through (7) and thereby eliminate the unobservable contract wages and expectations. This reduced form can be obtained as follows (recall that p_t is replaced by $\hat{p}_t$ in (5)): From (4) and (7) the conditional expectations of y_t and y_{t+1} are given by

$$\hat{y}_t = -0.5\beta(\hat{x}_t + x_{t-1}) + \theta\eta_{t-1}, \tag{8}$$

$$\hat{y}_{t+1} = -0.5\beta(\hat{x}_{t+1} + \hat{x}_t), \tag{9}$$

which can be substituted into (3) to obtain

$$\begin{aligned} x_t = {} & 0.5(x_{t-1} + \hat{x}_{t+1}) - 0.25\gamma\beta(\hat{x}_{t+1} + 2\hat{x}_t + x_{t-1}) \\ & + 0.5\gamma\theta\eta_{t-1} + u_t + \delta u_{t-1}. \end{aligned} \tag{10}$$

Eq. (10) describes the behavior of x_t in terms of the predetermined variables and the shock u_t. Taking conditional expectations in (10) given information through period $t-1$ we have

$$-0.5\hat{x}_{t-1} + c\hat{x}_t - 0.5\hat{x}_{t+1} = (1 - 0.5\gamma\beta)^{-1}(0.5\gamma\theta\hat{\eta}_{t-1} + \delta\hat{u}_{t-1}), \tag{11}$$

where

$$c = (1 + 0.5\gamma\beta)(1 - 0.5\gamma\beta)^{-1}. \tag{12}$$

Using the notation $L^s\hat{x}_r = \hat{x}_{r-s}$, eq. (11) can be written as

$$B(L)\hat{x}_t = (1 - 0.5\gamma\beta)^{-1}(0.5\gamma\theta\hat{\eta}_{t-1} + \delta\hat{u}_{t-1}), \tag{13}$$

where $B(L) = -0.5L + c - 0.5L^{-1}$. Factoring $B(L)$ into $\lambda A(L)A(L^{-1})$, where $A(L) = 1 - aL$, and dividing (13) through by $\lambda A(L^{-1})$ we obtain

$$A(L)\hat{x}_t = \lambda^{-1}A^{-1}(L^{-1})(1 - 0.5\gamma\beta)^{-1}(0.5\gamma\theta\hat{\eta}_{t-1} + \delta\hat{u}_{t-1}).$$

Since the difference between $\hat{x}_t$ and x_t is u_t the contract equation is therefore

$$x_t = ax_{t-1} + 2\pi_1\eta_{t-1} + \pi_2 u_{t-1} + u_t, \tag{14}$$

[9]Note that one might want both p_t and $\hat{p}_t$ on the right-hand side of (5). The conditional expectation would come from the monetary rule and the actual variable from the demand for money.

where

$$\pi_1 = 0.5\gamma\theta a(1-0.5\gamma\beta)^{-1}, \tag{15}$$

$$\pi_2 = 2\delta a(1-0.5\gamma\beta)^{-1}, \tag{16}$$

$$0 = 0.5 - ca + 0.5a^2. \tag{17}$$

The identities (15) and (16) define π_1 and π_2 as functions of the structural parameters γ, β, θ, δ, and the reduced form parameter a. To determine the parameter a, note that if $|c|>1$, eq. (17) has one solution with $|a|>1$ and one solution with $|a|<1$. Hence, there is a unique stable solution for a if $|c|>1$. This solution is given by

$$\begin{aligned} a &= c - \sqrt{c^2-1} \quad \text{if} \quad c>1, \\ &= c + \sqrt{c^2-1} \quad \text{if} \quad c<-1. \end{aligned} \tag{18}$$

The equation for the price level p_t is obtained by lagging eq. (14) one period, adding it to (14) and dividing by 2,

$$\begin{aligned} p_t &= ap_{t-1} + \pi_1(\eta_{t-1} + \eta_{t-2}) + 0.5u_t \\ &\quad + 0.5(1+\pi_2)u_{t-1} + 0.5\pi_2 u_{t-2}. \end{aligned} \tag{19}$$

An equation for output follows directly from (5) and when combined with (19) we have the vector representation

$$\begin{aligned} \begin{pmatrix} y_t \\ p_t \end{pmatrix} &= \begin{pmatrix} 0 & -\beta a \\ 0 & a \end{pmatrix} \begin{pmatrix} y_{t-1} \\ p_{t-1} \end{pmatrix} + \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix} + \begin{pmatrix} \theta - \beta\pi_1 & -\beta(1+\pi_2) \\ \pi_1 & 1+\pi_2 \end{pmatrix} \begin{pmatrix} u_{1t-1} \\ u_{2t-1} \end{pmatrix} \\ &\quad + \begin{pmatrix} -\beta\pi_1 & -\beta\pi_2 \\ \pi_1 & \pi_2 \end{pmatrix} \begin{pmatrix} u_{1t-2} \\ u_{2t-2} \end{pmatrix}, \end{aligned} \tag{20}$$

where we have set $u_{1t} = \eta_t$ and $u_{2t} = 0.5u_t$. This vector ARMA(1,2) model with constraints (15), (16), and (17) represents the basic reduced form of the model in terms of observable variables.

In Taylor (1980) this technique of polynomial factorization was used to solve the model in the case of N period contracts, for any $N \geqq 2$. For $N=2$ the factorization is trivial and simply requires the solution of the quadratic eq. (17). For larger N the factorization is more cumbersome computationally, but can be used in an analogous way for estimation and policy analysis.

3.1. Empirical regularities versus the theoretical model

One of our purposes is to compare the statistical properties of this rudimentary model with the empirical regularities presented in section 2 for each of the 10 countries. To do this we write the ARMA(1,2) model (20) in terms of its infinite moving average representation. Consider first the case where there is no exogenous serial correlation in the wage formation equation (so that all correlation in prices is due to the contracts), but where velocity has the correlation structure of eq. (7). Using the ψ_i notation of eq. (2) the theoretical moving average representation is then given by

Lag	ψ_{11}	ψ_{12}	ψ_{21}	ψ_{22}
0	1	0	0	1
1	$\theta-\beta\pi_1$	$-\beta(1+a)$	π_1	$1+a$
2	$-\beta\pi_1(1+a)$	$-\beta a(1+a)$	$\pi_1(1+a)$	$a(1+a)$
$i\geqq 3$	$-\beta\pi_i a^{i-2}(1+a)$	$-\beta a^{i-1}(1+a)$	$\pi_1 a^{i-2}(1+a)$	$a^{i-1}(1+a)$

The coefficient a is positive and less than one if $0<\beta\gamma<2$ which will be the case if $\gamma>0$ and $\beta>0$ (but not too large); that is, if wages respond positively to excess demand for labor, and aggregate demand is not too restrictive. A completely accommodative policy with $\beta=0$ would not satisfy this inequality. The coefficient π_1 will also be positive under these assumptions.

If β and γ lie in these parameter regions, the pattern of lag weights implied by the theoretical model has the same general features as the estimated parameters. In particular ψ_{21} and ψ_{22} are positive and begin to rise before tapering off to zero, with the shape of ψ_{21} tending to lag behind ψ_{22}. The ψ_{12} weights are negative and, if $\theta>\beta\pi_1$, the ψ_{11} weights are first positive before turning negative and then tapering off toward zero. Thus, price shocks have a strong positive persistence effect on prices and a negative cross effect on output, while real output shocks have a much shorter persistence effect on output and a positive cross-effect on prices. Moreover, the cross effects follow the same general shape (in opposite directions) and correspond to the shape of the price persistence. Finally, a univariate moving average representation for output would show stronger persistence effects than this bivariate representation.

It is important to note that all these lag weights depend on the policy parameter β. Larger values of β will tend to reduce a and thereby lower the persistence of prices. Larger β values will also strengthen the cross effects from prices to output (ψ_{12}), while very low values of β which characterize very accommodative policies will weaken the cross effects. In the extreme case of $\beta=0$, output is exogenous to prices. Hence, very accommodative policies will tend to generate reduced form equations in which real variables appear to be exogenous to nominal variables.

The *negative* effect of price shocks on output does not limit the ability of this model to generate a Phillips curve, that is, a *positive* correlation between inflation and output deviations from trend. As shown in Taylor (1980), the theoretical correlation between $p_{t+1}-p_t$ and y_t can be positive. After a positive price shock, for example, the rate of inflation declines relative to its long-run trend rate in order that the path of prices can return, at least partially, to the previous trend. During this period of less than normal inflation, output tends to be below normal because it was pushed down by the price shock and is slow to return. The reverse chain of events occurs after a negative price shock. As mentioned above, these types of correlations are found in the data. Table 1 shows that the correlation between $p_{t+1}-p_t$ and y_t is positive for all countries except Italy, even though p_t and y_t are negatively correlated.

In sum, this relatively simple macroeconomic structure has properties which correspond fairly closely with the major empirical regularities found in the unstructured statistical estimates for most of the countries. To be sure there are some discrepancies where persistence of output and prices is different from that implied by the model. Some of these discrepancies might simply reflect different contract lengths across the countries, or other forms of serial correlation. For example, letting δ vary from zero across countries can account for some of the differences. But, in general, the discrepancies point to the need for more investigation and perhaps additions or modifications to the rudimentary model.

3.2. The policy tradeoff

To show that this model implies a tradeoff between σ_y and σ_p, consider first the special case of $\theta=0$ and $\delta=0$. Then, from eq. (20) we have that

$$\sigma_p^2=2\sigma_2^2/(1-a), \tag{21}$$

$$\sigma_y^2=\beta^2[(1+a)/(1-a)]\sigma_2^2+\sigma_1^2, \tag{22}$$

where $\sigma_1^2=Eu_{1t}^2$ and $\sigma_2^2=Eu_{2t}^2$. As β increases, the parameter a decreases, and consequently σ_p^2 declines. At the same time, σ_y^2 rises. The graph of σ_p^2 versus σ_y^2 defines the tradeoff.

When $\theta\neq0$ and $\delta\neq0$ and these parameters cannot be altered by additional feedback policy instruments, the expressions for σ_y and σ_p are more complicated but the tradeoff can be computed numerically by inverting (20) into its pure moving average representation

$$z_t=\sum_{i=0}^{\infty}\Psi_i u_{t-i}, \tag{23}$$

and evaluating the variance–covariance matrix of z_t as

$$Ez_t z_t' = \sum_{i=0}^{\infty} \Psi_i \Omega \Psi_i', \tag{24}$$

where $\Omega = Eu_t u_t'$. The values of σ_y and σ_p can then be obtained from the diagonal elements of $Ez_t z_t'$. Since the moving average parameters Ψ_i depend on β in a known way, the (σ_y, σ_p) pair will also depend on β. Hence, the tradeoff can be computed just as in the more simple case $\theta = \delta = 0$. Clearly this method generalizes to higher order systems.

If it were thought that more complex feedback policy could reduce the serial correlation of velocity shocks, then the policy analysis would involve more than one policy parameter. In this case, the evaluation of the tradeoff requires computing optimal control policies for different objective function weights. This optimization ensures that the policy rules are 'efficient' and trace out an efficiency locus of σ_y and σ_p pairs corresponding to the tradeoff curve [see Taylor (1979)]. If we characterize the policy rule in terms of a single parameter β and assume that further feedback coefficients cannot offset the serial correlation in velocity (corresponding to the parameter θ), then we do not need to optimize over a set of policy parameters; the tradeoff curve can be traced out by varying β and observing the changes in σ_y and σ_p. In effect we take θ and δ as given and assume that policy has already reduced the serial correlation as much as possible. This means that we have defined the tradeoff so that each country is on the tradeoff curve and the policy problem is to choose a point on the curve. This approach can be interpreted as including in the economic structure any policy mistakes which are inherent or unavoidable. When we take the structure as given we implicitly assume that the economy is at its frontier.

4. Comparison of policies and structures

The preceeding comparison of the theoretical moving average coefficients with those estimated in the different countries indicates that a contracting model of the type introduced in section 3 is capable of generating some of the major empirical features of the aggregate time series. Such contracting models have the property that the way in which aggregate demand policy responds to the state of the economy matters for the behavior of output and that a tradeoff exists between the goals of price stability and output stability. Hence, this comparison between the model and the data lends support for the view that these policy effects and tradeoffs exist. Further support, which is less dependent on this particular model, comes from the rejection of output-exogeneity as reported in section 2.

It is important, however, to check whether the policy tradeoffs are stable in the sense that the structural parameters of the model are invariant to changes in the policy rule. A simple test of this invariance property would be to estimate the policy parameters and the structural parameters and examine whether there is any systematic relationship across the different countries. This requires estimating the parameters of a structural econometric model. While such a test might best be made within the framework of a very disaggregated model which is able to reflect the particular economic structure of each country, useful information is also revealed by tests based on a simple aggregate model, such as the one presented here. Later refinements can then serve as a check on the robustness of the results from the aggregate test.

4.1. Parameter estimation

From the reduced form of the model given in eq. (20) we can estimate the structural parameters β, γ, θ, and δ and the covariance matrix of the shocks u_{1t} and u_{2t} using maximum likelihood techniques. This reduced form is a vector autoregressive-moving average model [ARMA (1,2)] with non-linear constraints on the parameters. If we assume that the vector $u_t=(u_{1t}, u_{2t})'$ has a bivariate normal distribution, then we can compute the likelihood function in terms of the observations and the parameters, and hence in principle we can find the maximum of the likelihood function with respect to the parameters. This maximum likelihood procedure is equivalent to minimizing the determinant of $\sum_{t=1}^{T} u_t u_t'$, where $u_t=(u_{1t}, u_{2t})'$ with respect to β, γ, θ and δ, subject to the constraints. Although this determinant is a highly non-linear function of the structural parameters, the evaluation of the function is not particularly costly and conventional non-linear minimization routines can be used to find the minimum.[10] Sargent (1978) uses a similar approach for labor demand estimation.

The maximum likelihood estimates of the policy parameter β and the structural parameters γ, θ, and δ are shown in table 7. The serial correlation parameters θ and δ are significant for most of the countries, so that not all of the dynamic properties of the linearly detrended data can be explained by two-period wage contracts. This feature is also evident from the analysis of section 3. In order to determine whether the aggregate demand parameter β and the contract adjustment parameter γ are sensitive to the wage shock serial correlation assumption, the model was also estimated constraining δ to

[10]The non-linear minimum distance technique used by Taylor (1979) to estimate a rational expectations model for quarterly data in the U.S. required a grid search over the moving average parameters. Here the numerical technique maximizes the likelihood function jointly with respect to autoregressive and the moving average parameters. As in Taylor (1979) we maximize conditional on the initial values of $u_t=0$. The GRADX algorithm of the Goldfeld–Quandt program GQOPT was used for this maximization.

Table 7

Policy and structural parameter estimates, 1956–76.[a]

	β	γ	θ	δ	$L(4)$	$L(7)$	$L(12)$
Austria	0.0114	3.517*	0.788*	−0.373*	97.3	105.1	136.6
Canada	0.0901*	2.696*	0.782*	0.521*	106.6	117.1	123.8
Denmark	0.0373	0.609	0.868*	0.080	110.9	115.0	125.8
Germany	0.3727*	12.692*	0.714*	2.024*	109.2	117.1	129.5
Italy	0.2967*	0.315*	0.131	0.318*	103.8	107.4	128.5
Netherlands	0.0008	0.821	0.681*	−0.051	110.2	111.1	120.3
Norway	0.1255*	6.838*	0.714*	0.486*	112.8	122.3	128.5
Sweden	0.1317*	5.688*	0.901*	0.602*	118.0	123.5	140.7
United Kingdom	0.1165*	2.500*	0.351	0.688*	109.2	110.1	126.0
United States	0.2936*	0.087	0.549*	0.183	118.3	120.7	134.5

[a]*Note*: The structural estimates for β, γ, θ, and δ are obtained by estimating the reduced form (20) subject to the rational expectations constraints. The L columns give the value of the log of the likelihood function at the maximum. $L(4)$ corresponds to this 4 parameter model, $L(7)$ corresponds to a less restricted model described in the text, and $L(12)$ corresponds to the fully unrestricted ARMA(1,2) model. The asterisk indicates that the estimated coefficient is significantly different from zero at the 5 percent level.

Table 8

Policy and structural parameter estimates, 1956–76.[a]

	β	γ	θ	$L(3)$
Austria	0.0357	9.304*	0.753*	94.2
Canada	0.0746	1.768	0.845*	104.0
Denmark	0.0368	0.533	0.895*	110.6
Germany	0.3397*	5.888*	0.759*	105.5
Italy	0.3186*	0.129	0.164	101.8
Netherlands	0.0010	0.636	0.702*	110.0
Norway	0.0970	5.900	0.595	110.4
Sweden	0.1004	2.640*	0.886*	115.8
United Kingdom	0.0742*	0.925	0.491*	105.2
United States	0.2915*	0.046	0.550*	116.7

[a]*Note*: The structural estimates for β, γ, and θ are obtained by estimating the reduced form in the text subject to the rational expectations constraints and $\delta=0$. The column $L(3)$ gives the value of the log of the likelihood function at the maximum. The asterisk indicates that the estimated coefficient is significantly different from zero at the 5 percent level.

equal 0. The results of this estimation are reported in table 8, where it is clear that γ is more sensitive to the serial correlation assumption than β. Apparently constraining δ to 0 forces the wage contracting to absorb more of the persistence of wages by lowering the contract adjustment parameter. The effect is analogous to the influence of serial correlation correction in a simple regression equation with a lagged dependent variable. It is encouraging that β is not sensitive to this correction. For the purposes of intercounty comparisons we will focus on the results in table 7.

The ARMA(1, 2) model was also estimated without any constraints on the parameters, thereby increasing the number of parameters from 4 to 12. The unconstrained estimates gave a value for the likelihood function which was significantly greater than the constrained maximum in each of the countries.[11] In order to find the source for this significant difference we also estimated the model by letting the 5 parameters of the price equation vary freely but keeping the cross equation constraints intact. This results in a 7 parameter model [5 from the price equation plus β and θ as in (20)]. Likelihood ratio tests of the basic 4 parameter model against the 7 parameter model would then be a test of the dynamic constraints imposed on the price equation by the wage contracts and the rational expectations, as distinct from the cross equation constraints imposed by the aggregate demand policy rule. Tests of the 7 parameter model against the 12 parameter model would be a test of the cross equation constraints. For five of the countries – Denmark, Italy, the Netherlands, the U.K. and the U.S. – we could not reject the 4 parameter model against the 7 parameter model at the 5 percent significance level. However, the likelihood ratio tests indicate significant rejection of the 7 parameter model against the 12 parameter model for all countries at the 5 percent level. This suggests that the cross equation constraints, rather than the dynamic constraints, are restricting the likelihood function more significantly in these countries. This is similar to a result found by Barro (1978) for his model based on unanticipated monetary growth in the U.S.

4.2. *The relationship between β and γ across countries*

Using the terminology of section 3, all of the countries use a policy of partial accommodation ($0<\beta<1$), and all are far from the 'point for point' rule $\beta=1$, and even further from the 'Keynes' rule $\beta>1$. Several of the countries have β values very close to zero, in particular the Netherlands. Germany has the highest value of β which indicates its policy rule was the least accommodative of the countries examined in this paper. The United States appears to be more accommodative than Germany, and the United Kingdom more accommodative than both. This ranking of these three countries might suggest a policy tradeoff in the sense that Germany's output variance is highest and price variance is lowest of these three countries, while

[11]These results are based on likelihood ratio tests using the asymptotic χ^2 distribution with degrees of freedom equal to the number of constraints. Some of the unconstrained ARMA models were in the non-invertibility region which may indicate that our conditional likelihood function is too low which biases the tests toward rejection. An estimation procedure to constrain the model to lie in the invertibility region would be preferable.

the U.K. has the lowest output variance and the highest price variance. However, the fact that Italy has a high value for β indicates that such conclusions cannot be made without taking account of structural differences across the countries.

The variation of the structural parameter γ across the countries is very large, and this may raise some concern about our approximation that contract lengths are the same in each country. The low values of γ may simply reflect longer and more staggered contracts rather than different contract adjustment speeds. The parameter γ is extremely low for Denmark, Italy, the Netherlands, and the U.S., and relatively large for Austria, Germany, Norway and Sweden.

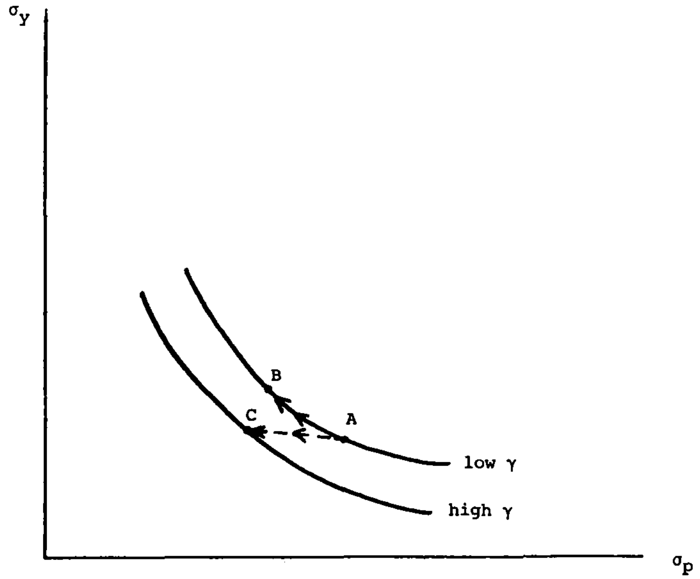

Fig. 1. Effect of a policy induced shift in the output–price stability tradeoff.

Any systematic relationship between these estimates of β and γ would raise doubts about the stability of the policy tradeoff. For example, if β and γ showed a systematic positive relationship, then one might suspect that a policy of less accommodation would be accompanied by greater wage flexibility (that is, a higher β might cause γ to rise). Since greater wage flexibility would enable the economy to operate with smaller business cycle fluctuations, the parameter shift would tend to move the tradeoff and offset the real effects of the more restrictive policy. This possibility is illustrated in fig. 1. Starting at point A policymakers move toward a more restrictive policy. If the tradeoff is stable, then the result is larger business cycle fluctuations. However, if wages become more flexible, then the tradeoff shifts down and the increase in business cycle amplitude is reduced, or perhaps offset, as at point *C*. From a policy perspective, it is therefore important to check whether γ might move in this direction.[12]

The estimated values for γ and β presented in table 7, indicate a slight positive relationship between these two parameters. Clearly it would not be wise to rule out the possibility of a relationship between γ and β based on this evidence. However, the high values of β and γ in Germany tend to dominate the evidence and without this sample point the relationship would not be significant. Note that Italy and the U.S., which have the second and third most restrictive policies have very sluggish adjustment parameters. In sum, the evidence is mixed, but certainly does not give the stable tradeoff view any additional support over that presented in section 3.

A word of caution is necessary in interpreting the observed relationship between policy and the structural parameters, however. In using a cross-section of parameter estimates to test such relationships, one usually assumes that the direction of causation, if any, goes from policy to structure. In this case, the degree of monetary accommodation is suspected of influencing adjustment speeds. It is important to note, however, that the relationship might go the other way. If policy-makers perceived that the economy is constrained by very sluggish wage adjustments, then they might be reluctant to implement restrictive policies; doing so would increase the amplitude of business cycles by more than would be desirable. On the other hand, if the economy had a high value of γ, then implementing a high β policy would increase business cycle fluctuations only slightly. Hence, even if γ were not affected by changes in β, one would expect to see a positive correlation between β and γ. For testing the sensitivity of structure to policy, therefore, a simultaneous equations framework would be required. Obtaining identification of the two way relationship would be quite difficult even if the sample of countries was large enough. But unless social preferences for output stability versus price stability are perverse, this simultaneous equations problem would tend to 'bias' the estimates toward a significant positive relationship.[13]

[12]It is interesting to note that an increase in wage adjustment speeds in response to a more inflation-conscious policy would tend to make the reduced form Phillips curve *steeper*. Lucas (1973) tested for the opposite effect and found evidence for it by comparing the flat Phillips curve in relatively stable inflation countries with the steep Phillips curves in unstable inflation countries such as Argentina and Paraguay.

[13]The estimated covariance between the estimated β and γ is *negative* for all countries except Austria and Sweden. Hence, sampling variability is unlikely to cause a positive relationship.

5. Concluding remarks

The aim of this international comparison has been to examine the tradeoff between aggregate output and price stability which is implied by contracting models with rational expectations. Much of the evidence is consistent with the existence of such a tradeoff, although certain discrepancies suggest the need for further research.

Such research might include econometric extensions such as an examination of alternative detrending procedures and serial correlation corrections, or simply the use of quarterly data and information from additional countries. There is also a need for theoretical research on the determinants of contract lengths and the degree of staggering, the channels of transmission of shocks across countries, and, perhaps most important, the simultaneous relationship between economic policy and economic structure. The experience with the estimation and policy analysis techniques presented here, indicates that they should prove useful for many of these suggested lines of research.

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