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IS P-STAR A GOOD INDICATOR OF INFLATIONARY PRESSURE IN OECD COUNTRIES?

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CONTENTS

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INTRODUCTION

The P-star (P*) concept was first developed by the U.S. Federal Reserve as a simple, yet comprehensive, indicator of inflationary pressure (Hallman et al., **1989)**[1]. P* is defined as the price level which is consistent with current money supply and equilibrium in goods and financial markets. As the gap between the actual price level (P) and P* is zero in equilibrium, deviations of P from P* indicate the amount of price adjustment which has not yet materialised and can help predict future movements in the price level. In all standard models of inflation the output gap is a major explanatory variable for inflation. The important novelty of the P* approach, however, is that deviations of the velocity of money from "trend" levels also matter for price-level determination.

The P* approach is not new: it is based on the assumption that in the long run the price level is determined by money supply, following the classic tradition of the quantity theory of the money. One can find a variant of the P* approach as far back as David Hume (see Humphrey, **1989** for a review of the precursors). The validity of money-supply-driven explanations of the price level hinges on the existence of an identifiable trend in the income velocity of money and the assumption that potential output is not affected by monetary policy[2]. The relevance of the P* approach for policy implementation also depends on the ability of the monetary authorities to influence monetary aggregates in the short and long runs (Pecchenino and Rasche, **1990).** Furthermore, for the P* concept to be truly useful, it must provide information not captured by other inflation models (Christiano, **1989;** Haslag, **1990;** Kuttner, **1990).**

To date, the approach has only been tested in a few countries. The purpose of this paper is to assess its usefulness for a large sample of OECD countries. The paper is organised as follows. Section I describes the theoretical framework. Section II presents the data and estimates for the potential price level, P*. This is followed by a section which reports regression results for 20 OECD countries. The in-sample tracking ability and forecasting performance of the P* equations are then compared to rival inflation models and to official OECD projections. Conclusions are presented in a final section.

I. THE P* APPROACH

Calculation of P* takes the quantity theory of money as a starting point. Let P be the actual price level, M a (nominal) monetary aggregate, V the income velocity of money and Q output at constant prices. The velocity identity is:

$$P \equiv M \times V/Q \qquad [1]$$

Denoting equilibrium (trend) values by "*", the identity becomes:

$$P^* \equiv M \times V^*/Q^* \qquad [2]$$

At trend velocity, P* is proportional to the money stock per unit of potential output. Dividing the second by the first equation and taking logs (lower-case notation) gives:

$$p^* - p \equiv (v^* - v) + (q - q^*) \qquad [3]$$

From the gap between equilibrium and actual prices, $p^* - p$, the P* model predicts the direction of movement of the price level: it will rise, fall or remain unchanged as the actual price level is below, above or at the equilibrium level. The price gap, however, does not contain information about the dynamics of adjustment of P to P*. In this paper, an error-correction model of the adjustment process was adopted. A general dynamic specification of such a model is:

$$dp = a\,(p^*_{-1} - p_{-1}) + \sum_{i=1}^{n} b_i\, dp_{-i} + \sum_{j=0}^{m} c_j\, dp^*_{-j} \qquad [4]$$

The coefficient a, the speed of adjustment of prices to P*, should be positive and the coefficients b and c sum to one in order to ensure the equality of the actual and potential price levels in the long run.

In equation [4], the two components of the price gap, the velocity and output gaps, are constrained to have the same coefficient. Considering both gaps separately gives the following equation:

$$dp = a_1\,(v^*_{-1} - v_{-1}) + a_2\,(q_{-1} - q^*_{-1}) + \sum_{i=1}^{n} b_{1i}\, dp_{-i} + \qquad [5]$$

$$\sum_{j=1}^{m} c_{1j}\,(dv^*_{-j} - dv_{-j}) + \sum_{j=1}^{m} c_{2j}\,(dq_{-j} - dq^*_{-j});$$

$$a_1,\ a_2,\ \sum c_{1j},\ \sum c_{2j} > 0 \text{ and } \sum b_{1i} = 1$$

New-Keynesian approaches towards modelling inflation focus on the output gap and inertia in price adjustment. In these models, inflation increases when the output gap opens up, while the price level is indeterminate ($a_1 = 0$, $a_2 > 0$, $\sum c_{1j} = 0$). Gordon (1990) has extended this approach by also including an output gap in first differences as a proxy for possible hysteresis effects:

$$dp = a_2\,(q_{-1} - q^*_{-1}) + \sum_{j=1}^{m} c_{2j}\,(dq_{-j} - dq^*_{-j}) + dp^e \qquad [6]$$

In such an inflation model, expectations (dp^e) are assumed to be adaptive and represented by past inflation. The new-Keynesian inflation model and the price-gap model coincide if the set of information available to agents is assumed to include velocity gap terms[3]:

$$dp^e = \sum_{i=1}^{n} b_{1i}\, dp_{-i} + a_1\,(v^*_{-1} - v_{-1}) + \sum_{j=1}^{m} c_{1j}\,(dv^*_{-j} - dv_{-j}) \qquad [7]$$

The usefulness of the price gap as an indicator of inflationary pressure must be qualified in two respects:

i) In a fixed exchange-rate regime, money supply becomes endogenous. Monetary developments in the reserve-currency country may be a better indicator of inflationary pressure than the domestic money stock (Browne, **1986).** Thus, for instance, money supply in Germany might be more relevant for explaining inflation in other EMS countries than domestic money supply, as the domestic money stock could lag price developments.

ii) As the equation focuses on the adjustment towards long-run equilibrium levels, it does not capture important factors influencing prices in the short run, e.g. indirect tax changes, food or energy price shocks. P* then provides a measure of where the price level will go after such transitory shocks have worked themselves out.

II. CALCULATION OF P*

As implied by the basic identities, the GDP deflator is used as the price variable, and real GDP as the output variable. Different money aggregates exist for each country and there is little guidance as to which of the available time-series corresponds best to the theoretical "money" concept. For the purposes of this paper, we have used the money aggregates projected during OECD's half-yearly forecasting exercises; these aggregates are either targeted or closely monitored by the monetary authorities in the countries in question. For the large countries, M2 is used for the United States, Italy and Canada, M2 plus certificates of deposits for Japan, M3 for Germany and France[4] and M4 for the United Kingdom.

A key question in implementing the P* approach is how to measure potential output (Q*) and trend velocity (V*). Estimates of P* in this paper are based on **two** alternatives: *i)* simple linear time trends; and *ii)* stochastic trends of the real GDP and velocity series, using a filtering approach.

A number of studies assume that output and velocity follow a deterministic path in the long run. For instance, Christiano **(1989)** and Hannah and James **(1989)** used a linear time trend to calculate potential output. Trend velocity since **1954** is measured by its average value in the Federal Reserve's study and in Christiano **(1989),** while the Bank of Japan study employs a linear time trend.

Time-series analysis suggests, however, that real GDP in OECD countries contains a unit root, i.e. it follows a stochastic rather than deterministic trend[5]. The stationarity of velocity has been most often investigated using a cointegration framework. This approach tests for the existence of a stable long-run relationship between money, real income and prices. But the existence of a cointegrating vector between these variables is only a necessary condition for velocity to be stationary[6]. Unit-root tests – which are conceptually equivalent to cointegration tests if unitary elasticities of prices with respect to money and real income are imposed – are more appropriate in the context of the P* approach. Unit-root test statistics reported in Table 1 suggest that for most countries, including the United States, velocity did not tend to revert to some mean value or

Table 1. Unit root tests: logarithm of money velocity

	T	Start period	Alternative 3		Alternative 2	Alternative 1
			0 3	Φ2	O1	t
United States	(60)	S1 1960	4.01	2.68	4.21	–0.24
Japan	(46)	S1 1967	1.30	3.39		–1.72
Germany	(42)	S1 1969	3.03	5.81		
France	(40)	S1 1970	3.20	2.16	1.89	–0.38
Italy	(52)	S1 1964	2.77	1.85	1.05	–0.27
United Kingdom	(54)	S1 1963	1.20	1.35	0.93	–1.29
Canada	(44)	S1 1968	2.00	2.71	2.33	–1.99*
Australia	(59)	S2 1960	3.30	2.20	0.84	–0.09
Austria	(44)	S1 1968	7.46'			
Belgium/Luxembourg	(40)	S1 1970	1.12	1.32	1.03	–1.32
Denmark	(54)	S1 1963	2.10	1.41	2.46	–0.33
Finland	(60)	S1 1960				–1.23
Greece	(60)	S1 1960	4.56	6.98		
Ireland	(60)	S1 1960	5.56	3.86	1.93	0.42
Netherlands	(60)	S1 1960	3.53	2.72		-0.95
New Zealand	(48)	S1 1966	3.01	2.25	0.40	–0.84
Norway	(40)	S1 1970	2.82	1.94	2.61	–0.56
Spain	(42)	S1 1969	7.37*			
Sweden	(60)	S1 1960	3.27	2.83	1.00	1.28
Switzerland	(36)	S1 1972	7.24*			
OECD	(40)	S1 1970	1.24	3.01		–2.39'

In parentheses: number (T) of observations (semi-annual data). An asterisk denotes that the non-stationarity hypothesis can be rejected.

Description of the tests

The unit root tests reported here are augmented Dickey-Fuller tests with second-order correction. The testing strategy, following Perron (1988), involves a sequence of tests that runs from general to restricted alternative hypotheses. We begin with the alternative hypothesis of a stationary series with a time trend:

$$\Delta\chi_t = \mu + \beta(t - T/2) + \alpha\chi_{t-1} + \sum_{i=1}^{2} \gamma_i \Delta\chi_{t-1} + u_t \qquad [a1]$$

If $\beta = 0$, there is no time trend; if $a = 0$, the series has a unit root (that is, it is non-stationary). If the series has a unit root, μ is interpreted as its drift. The γ-coefficients are the second-order correction terms.

The test statistic Φ3 jointly tests the two zero restrictions ($a = \beta = 0$) for the null of a unit root, no time trend and a drift. If it exceeds its critical value, the null is rejected, i.e. the series is deemed to be stationary, and the process stops. If not, the null is respecified to have no drift and no time trend. The test statistic Φ2 therefore jointly tests three zero restrictions ($\mu = a = \beta = 0$). If it exceeds its critical value, the additional constraint of zero drift is rejected, the series is deemed to be non-stationary on the strength of Φ3, and the process ends. If not, further tests are carried out with more restricted alternatives. The first restriction yields an alternative of a stationary series with no time trend i.e. the same as [a1] but without the second term. The relevant null is a series with a unit root and no drift. Therefore, the statistic Φ1 jointly tests two zero restrictions ($p = a = 0$). If it exceeds its critical value, the series is deemed to be stationary. If not, a final restricted alternative – a non-stationary series with a zero mean and no time trend – is considered, i.e. the same as [a1] but without the first two terms. The statistic t tests one zero restriction ($a = 0$), for the null of a unit-root. If it exceeds its critical value, the series is deemed to be stationary. Thus, only if all four test statistics lie within their critical values, or if 0 3 is below and 0 2 is above its critical value, is the series deemed to have a unit root.

For 50 observations, the critical value for the t-statistic at the 5 per cent is – 1.95 (Fuller, 1976), *Introduction to Statistical* Time *Series*, p. 373). Critical values for Φ3, 0 2 and 0 1 are 6.73, 5.13, 4.86 (Dickey and Fuller (1981), Econometrica 4, p. 1063).

deterministic time trend over the sample periods'. The null hypothesis of non-stationarity could be rejected only in the cases of Canada, Austria, Spain, Switzerland and the OECD in aggregate[8].

The presence of unit roots in output and velocity implies that they do not revert to some deterministic time trends or historical averages in the long run. Thus, the use of time trends or mean values for calculating potential output and equilibrium velocity can yield non-stationary price gaps, which is inconsistent with the assumption of the P* model[9]. Therefore, alternative approaches should be used.

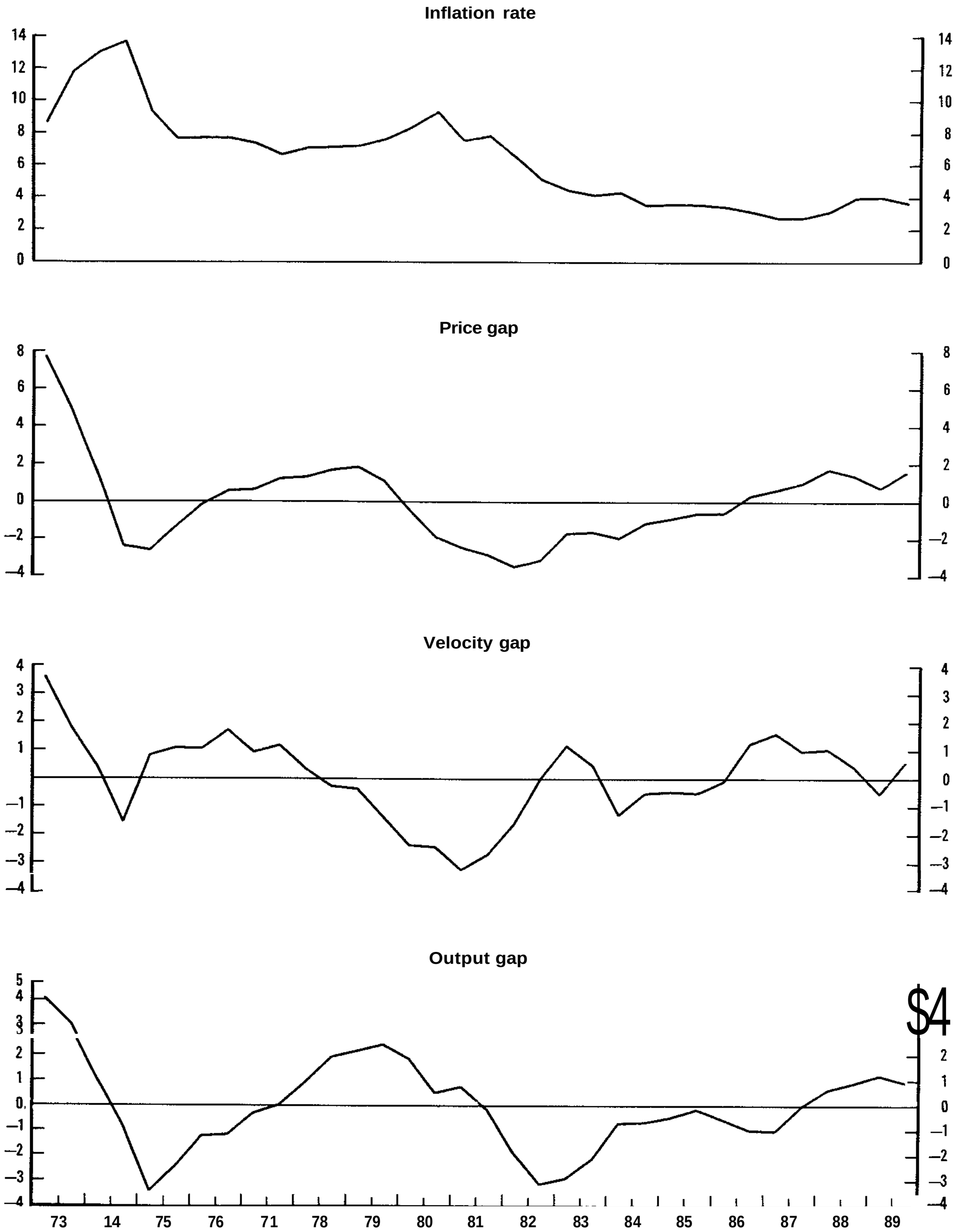
The first alternative is to use structural models of the determination of potential output and equilibrium velocity. For example, the Federal Reserve's studies, Ebrill and Fries **(1990),** and Pecchenino and Rasche **(1990),** use estimates of potential from Braun **(1990),** who derives them by combining a Phillips-curve-based estimate of the natural rate of unemployment with Okun's law. The Bank of Japan study uses an aggregate production function framework to derive its estimate of potential. As to the velocity of money, Hallman et *al*. **(1990)** use for the United States for the period **1870** to **1954** the fitted values from a regression of velocity on the fraction of the labour force employed in agriculture (a proxy for the industrialisation and of monetisation of the U.S. economy). Ebrill and Fries **(1990)** calculate the U.S. velocity gap as the residuals from a cointegrating equation explaining long-run velocity by the own and competing rates of return on M2.

A second alternative for computing potential output and equilibrium velocity is the use of filters. Using the Kalman filter, Bomhoff **(1990),** for instance, found that income velocity has not become less predictable in the **1980s** for the large OECD countries. For this paper, we prefer to use the Hodrick-Prescott filter[10], which is an appropriate filter for stochastic trends. The computational ease of this filter is a key advantage over the more complicated Kalman filter, especially when 20 countries are examined. Also the regression results reported below proved to be slightly better when the Hodrick-Prescott measure of trend output was used rather than a more sophisticated measure of potential derived from the supply block of the OECD's lnterlink model (Torres and Martin, **1990).**

The split of output into "transitory" and "trend" components with the Hodrick-Prescott filter depends on a smoothing factor which must be chosen by the user. A very small smoothing factor implies that most of the shocks to the series are changes in trend, while a very large factor leads to an almost constant trend so that virtually all shocks are transitory. There exists a critical degree of smoothing below which the resulting trend can preserve the long-run non-stationary properties of the series, while the deviations from the trend are made stationary (that is, only weakly auto-correlated). The smoothing factor has been selected on this basis. This choice of the smoothing factor ensures stationarity of the price gap, which is required by the P* approach.

Once potential output and trend velocity are defined, price gaps are calculated according to equation [3]. For illustrative purposes, the inflation rate is shown in the first panel of Chart **1** and the price gap ($p^* - p$) based on the Hodrick-Prescott filter in the second panel for the OECD as a whole[11]. Inflation tended to increase when the price gap became positive and tended to fall once the price gap became negative. Lags between changes in the sign of the price gap and inflation are between one and two years. Following a prolonged period of negative price gaps in the early to **mid-1980s** accompanied by falling inflation rates, the price gap became positive in

Chart 1. The price gap and its components for the OECD aggregate



Source: Data based on OECD, *Economic Outlook No. 47*

Chart 2. Actual and trend output
In logarithm

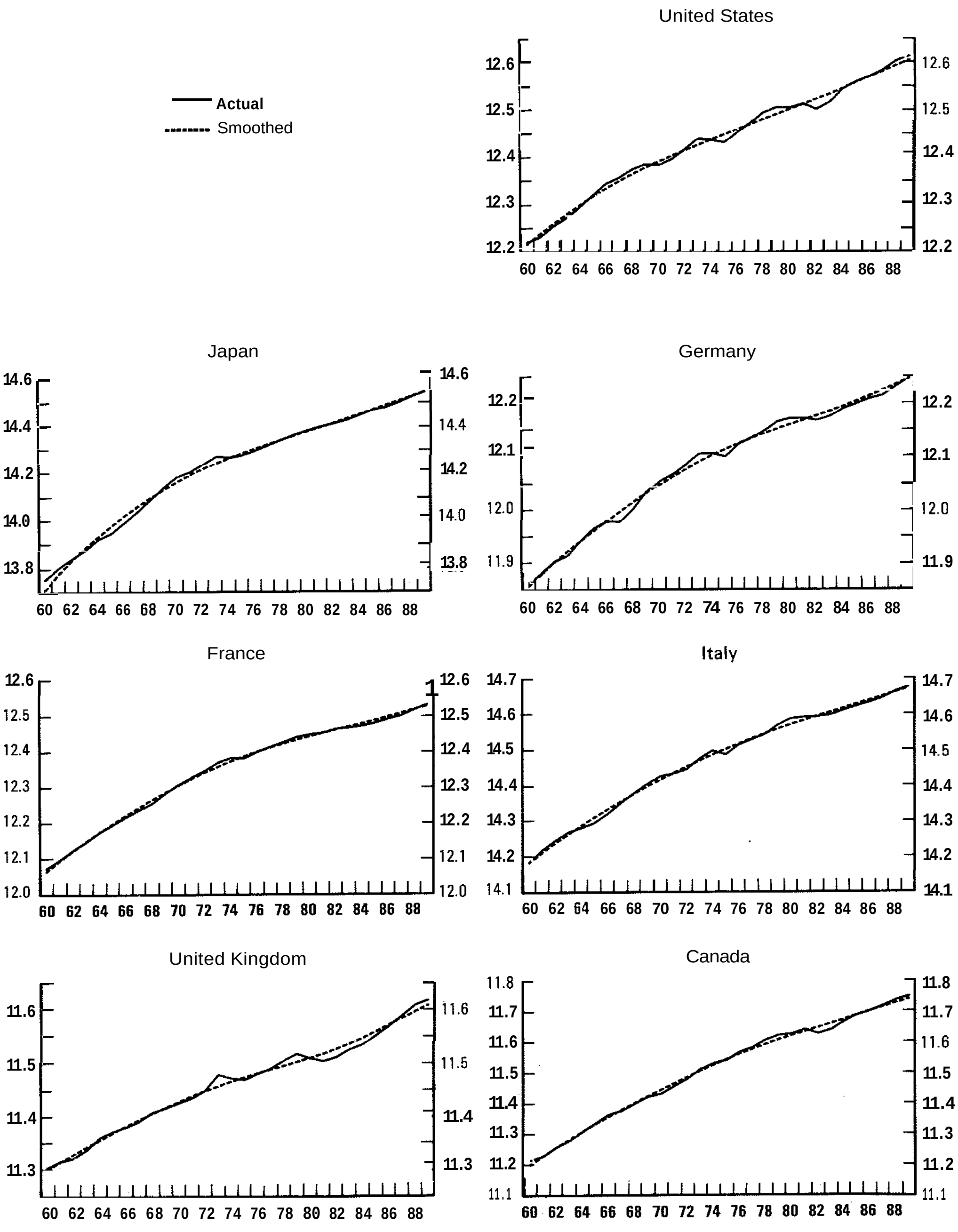
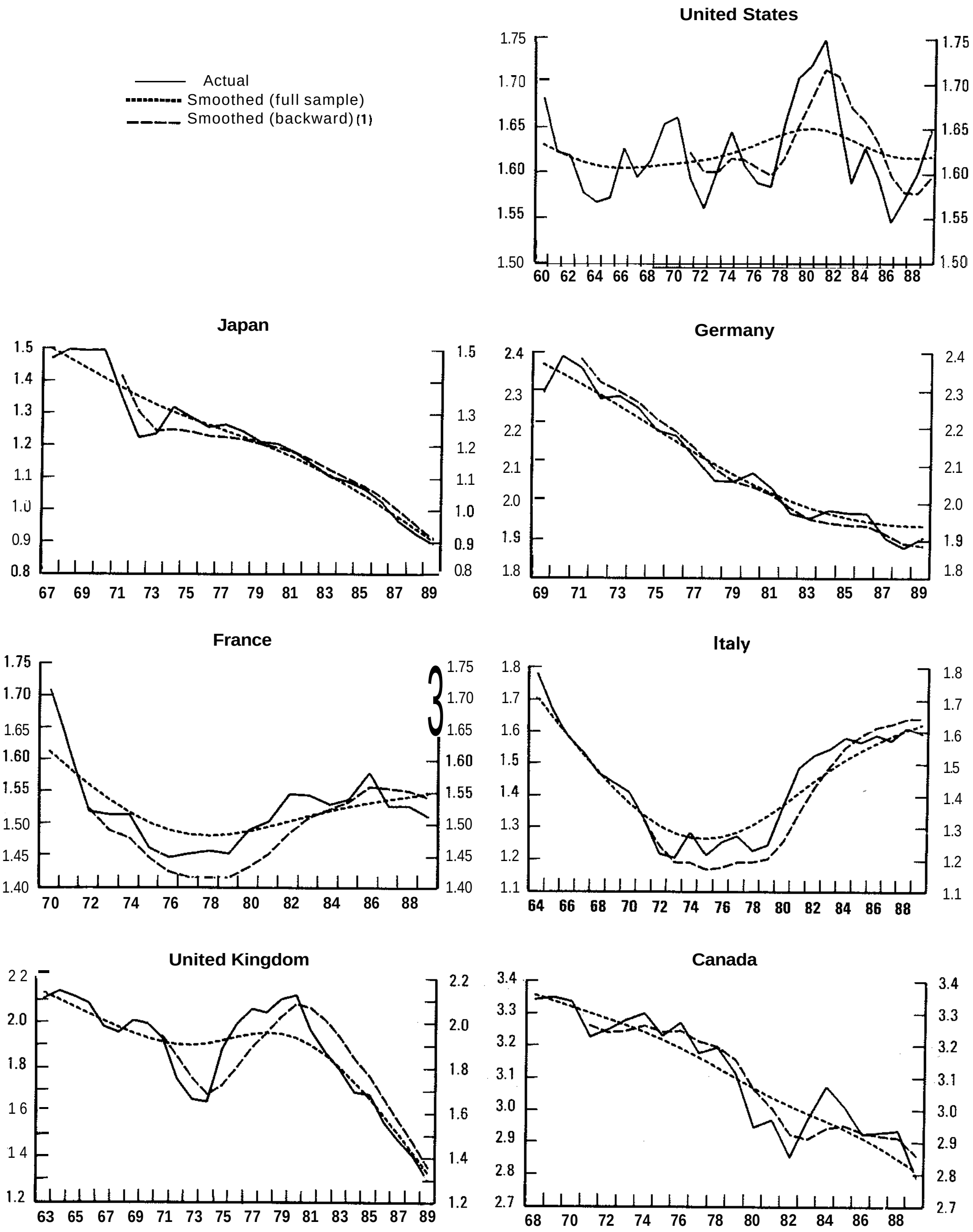


Chart 3. Actual arid trend velocity



1 No out-of-sample information is used The value of trend velocity in 1982 for the United States, for instance, is calculated by applying the filter to the data from 1960 to 1982

1987 and inflation has since drifted up again. The velocity gap is shown in the third panel and the output gap in the fourth panel of Chart **1.** Actual and trend output and velocity are shown for the seven large countries in Charts 2 and 3.

III. REGRESSION RESULTS

After experimentation with various lag distributions, the following error-correction specifications of the relationship between the actual and the potential price level were **selected**[12]:

$$dp = a_0 + a\,(p^*_{-1} - p_{-1}) + a_3\,dp_{-1} + a_4\,dp_{-2} + a_5\,dp_{-3} \qquad [8]$$

$$dp = a_0 + a_1\,(v^*_{-1} - v_{-1}) + a_2\,(q_{-1} - q^*_{-1}) + a_3\,dp_{-1} + a_4\,dp_{-2} + a_5\,dp_{-3} \qquad [9]$$

The dependent variable is the inflation rate. The first term on the right-hand side of equation **[8]** is the previous period price gap; in equation [9] this is split up into its two components. If the intercept, a_0, is zero and the coefficients on lagged inflation sum to unity, the actual and potential price levels will be equal in the **long-run**[13]. If, in addition, the constraint that $a_1 = a_2$ in equation **[9]** is accepted by the data, no forecasts of actual velocity and output are needed in order to forecast inflation; knowledge of trend velocity, potential output and future money-stock development is sufficient. Table 2 reports the results from estimating equation **[8]** and Table **3** from estimating equation **[9]**. Estimates are based on semi-annual data for the seven major countries and yearly data for the smaller countries (for these countries, a two-year lagged inflation term is used).

The results in Table 2 suggest that a satisfactory price-level equation can be estimated for most countries using the price gap and lagged inflation as explanatory variables. The price gap was not significant in the equations for Denmark, New Zealand and Norway, using either of the two methods for calculating trends. For most countries, the trend calculations using the Hodrick-Prescott filter improve the fit of the equations. Apart from the equation for Finland, however, the improvement is marginal, as the major explanation **of** inflation is its own past.

Table 3 presents results of equations with the separate components of the price gap as regressors. Equations using data generated by linear time trends are not shown in Table 3 except for the few countries for which the standard errors proved to be lower than those based on data generated by the Hodrick-Prescott filter. Except for Austria and Spain, the velocity gap based on the Hodrick-Prescott filter was found to be more significant than the one calculated using a linear time **trend**[14].

While the coefficient of the output gap was significant for all countries with the exception of the Netherlands, New Zealand and Norway, the coefficient of the velocity gap was insignificant for Germany, France, Australia, Austria, Belgium, Denmark, the Netherlands, New Zealand, Norway and Switzerland and wrongly signed in two of these countries[15]. Out of the 20 countries examined, the velocity gap has a correctly signed and significant coefficient for nine as well as for the OECD aggregate equation[16]. Note that only one monetary aggregate is tested in this paper. The explanatory power of other monetary aggregates and, hence, velocity gaps may be greater.

Table 2. Regression results: restricted P* model

Dependent variable: first difference of the log of the GDP deflator (Δln P)

		Intercept x 100	Price gap (–1)	Δln P (–1)	Δln P (–2)	Δln P (-3)	SEE x 100 adj. R^2	DW (h – stat.)	Start period[1]
United States	A.	0.16	0.07	0.69	0.16	0.07	0.58	2.08	S1
		(0.83)	(3.56)	(5.50)	(1.03)	(0.54)	0.78	(-1.44)	1962
	B.	–0.00	0.12	0.72	0.21	0.08	0.57	2.06	
		(-0.03)	(3.97)	(5.83)	(1.33)	(0.62)	0.79	(-0.88)	
Japan	A.	0.81	0.09	0.75	-0.32	0.02	0.97	2.10	s2
		(3.36)	(5.36)	(5.45)	(-1.84)	(0.15)	0.81	(–0.90)	1967
	B.	0.56	0.20	0.80	-0.27	0.20	0.94	2.13	
		(2.51)	(5.81)	(6.33)	(-1.62)	(1.77)	0.83	(-0.98)	
Germany	A.	0.84	0.10	0.27	0.42	-0.21	0.75	2.04	s2
		(2.60)	(2.90)	(1.83)	(2.98)	(-1.43)	0.50	(-0.52)	1969
	B.	0.37	0.19	0.32	0.55	–0.08	0.72	2.07	
		(1.22)	(3.45)	(2.23)	(4.15)	(-0.59)	0.53	(-0.66)	
France	A.	1.40	0.09	0.22	0.15	0.18	0.84	2.01	s2
		(2.80)	(3.65)	(1.43)	(1.01)	(1.28)	0.73	(-0.81)	1970
	B.	0.62	0.16	0.37	0.23	0.23	0.89	2.04	
		(1.40)	(2.94)	(2.44.)	(1.48)	(1.49)	0.70	(-1.38)	
Italy	A.	1.43	0.07	0.49	0.01	0.16	1.36	2.09	s2
		(3.21)	(4.39)	(3.71)	(0.09)	(1.34)	0.78	(-1.09)	1964
	B.	0.75	0.24	0.54	0.10	0.21	1.27	2.08	
		(2.C)	(5.40)	(4.50)	(0.72)	(1.84)	0.81	(-0.67)	
United Kingdom	A.	0.66	0.03	1.07	-0.39	0.19	1.45	2.03	S2
		(1.61)	(1.77)	(7.62)	(-1.96)	(1.31)	0.71	(0.73)	1963
	B.	0.69	0.11	0.98	-0.30	0.16	1.34	2.02	
		(1.89)	(3.40)	(7.40)	(-1.60)	(1.21)	0.75	(–1.01)	
Canada	A.	1.12	0.12	0.63	0.27	-0.11	0.87	1.85	S2
		(3.31)	(3.30)	(3.99)	(0.14)	(-0.79)	0.70	(0.73)	1968
	B.	1.18	0.17	0.67	0.04	-0.06	0.90	1.75	
		(3.01)	(2.70)	(4.13)	(0.18)	(-0.40)	0.68	(0.71)	
OECD total	A.	2.52	0.24	0.73	-0.20		0.87	1.76	1973
		(3.45)	(5.33)	(4.61)	(-1.37)		0.87	(0.68)	
	B.	1.12	0.46	0.91	-0.11		0.78	1.69	
		(2.01)	(6.17)	(6.97)	(-0.77)		0.89	(-0.18)	
EMS	A.	2.92	0.16	0.83	-0.34		0.73	2.11	1971
		(3.41)	(4.10)	(4.76)	(-2.10)		0.86	(-0.46)	
	B.	1.06	0.36	1.01	-0.19		0.67	2.17	
		(1.58)	(4.79)	(6.85)	(-1.22)		0.88	(-0.78)	
Australia	A.	1.58	0.05	1.13	-0.32		1.72	1.77	1963
		(2.23)	(1.01)	(5.81)	(-1.63)		0.79	(0.81)	
	B.	2.03	0.27	1.03	-0.27		1.59	1.68	
		(2.99)	(2.19)	(5.33)	(-1.55)	..	0.82	(3.11)	
Austria	A.	1.75	0.18	0.51	0.06	..	1.05	2.07	1969
		(2.27)	(3.21)	(2.50)	(0.29)	..	0.67	(-1.46)	
	B.	1.21	0.23	0.72	0.04	..	1.21	1.92	
		(1.34)	(1.93)	(3.29)	(0.17)	..	0.56	(0.63)	

Table 2. ***(Continued)*** **Regression results: restricted P* model**

Dependent variable: first difference of the log of the GDP deflator (Δln P)

		Intercept x 100	Price gap (−1)	Δln P (−1)	Δln P (−2)	SEE x 100 adj. R^2	DW (h − stat.)	Start period[1]
Belgium-Luxembourg	A.	3.19	0.13	0.81	−0.42	1.69	2.16	1971
		(2.89)	(2.30)	(3.66)	(−1.81)	0.59	(−1.60)	
	B.	2.45	0.27	0.98	-0.39	1.68	2.16	
		(2.36)	(2.36)	(4.56)	(−1.74)	0.59	(−1.20)	
Denmark	A.	1.20	0.03	0.79	0.03	1.70	1.77	1964
		(0.79)	(0.65)	(3.58)	(0.14)	0.48	(0.54)	
	B.	1.48	0.03	0.79	0.03	1.71	1.74	
		(1.03)	(0.38)	(3.38)	(0.14)	0.47	(0.53)	
Finland	A.	2.16	0.22	0.79	−0.01	2.76	1.92	1963
		(1.34)	(1.84)	(4.13)	(-0.06)	0.43	(0.70)	
	B.	1.98	0.68	0.79	-0.02	2.09	2.12	
		(1.78)	(4.77)	(5.47)	(-0.11)	0.67	(−1.15)	
Greece	A.	1.46	0.14	0.67	0.18	3.21	2.00	1963
		(1.19)	(2.80)	(3.66)	(1.02)	0.77	(-0.28)	
	B.	0.57	0.70	0.72	0.27	2.49	2.07	
		(0.59)	(5.32)	(5.16)	(1.93)	0.86	(-0.45)	
Ireland	A.	6.56	0.36	0.22	−0.00	2.94	2.03	1963
		(4.41)	(4.73)	(1.33)	(-0.05)	0.66	(-0.26)	
	B.	3.20	0.42	0.47	0.19	3.41	1.94	
		(2.15)	(3.29)	(2.73)	(1.15)	0.54	(0.04)	
Netherlands	A.	1.78	0.15	0.71	−0.01	1.75	1.58	1963
		(2.16)	(2.11)	(3.99)	(-0.07)	0.64	(1.92)	
	B.	1.28	0.13	0.77	-0.02	1.85	1.51	
		(1.56)	(1.19)	(4.18)	(−0.01)	0.60	(3.68)	
New Zealand	A.	6.98	−0.14	0.69	−0.35	3.62	1.99	1967
		(2.73)	(−1.03)	(3.24)	(−1.52)	0.37	(0.46)	
	B.	5.08	0.03	0.79	−0.23	3.72	1.96	
		(2.59)	(0.19)	(3.47)	(−1.08)	0.34	(0.43)	
Norway	A.	2.28	0.24	0.45	0.19	3.16	2.02	1971
		(0.94)	(1.56)	(1.99)	(0.77)	0.13	(−1.01)	
	B.	1.69	0.33	0.51	0.22	3.10	2.13	
		(0.67)	(1.74)	(2.21)	(0.89)	0.16	(−5.64)	
Spain	A.	3.41	0.20	0.38	0.26	1.75	2.23	1970
		(3.01)	(4.42)	(1.88)	(1.45)	0.84	(−1.29)	
	B.	3.02	0.33	0.62	0.12	2.09	1.76	
		(2.25)	(2.97)	(2.88)	(0.60)	0.77	(1.09)	
Sweden	A.	3.20	0.15	0.42	0.09	2.43	2.09	1963
		(2.59)	(2.45)	(2.23)	(0.52)	0.41	(−1.47)	
	B.	2.02	0.33	0.55	0.17	2.44	2.19	
		(1.64)	(2.37)	(2.94)	(0.93)	0.40	(−2.35)	
Switzerland	A.	2.13	0.14	0.60	-0.20	1.52	1.93	1973
		(2.66)	(2.26)	(2.74)	(−1.00)	0.50	(0.31)	
	B.	1.73	0.23	0.56	−0.06	1.38	2.11	
		(2.40)	(3.01)	(2.84)	(-0.30)	0.59	(-0.43)	

A. Price gap calculated using linear time trends.
B. Price gap calculated using the Hodrick Prescott filter.
1. End period = 1989.

Table 3. Regression results: unrestricted P* model

Dependent variable: first difference of the log of the GDP deflator (Δln P)

	Intercept x 100	Velocity gap (−1)	Output gap (−1)	Δln P (−1)	Δln P (−2)	Δln P (−3)	SEE x 100 adj. R^2	DW (h-stat.)	Homog. F-test[1]	Start period[2]
United States	0.12	0.08	0.18	0.59	0.20	0.17	0.55	2.08	0.22	S1
	(0.60)	(2.28)	(4.42)	(4.48)	(1.35)	(1.34)	0.80	(−2.68)	5.77	1962
Japan	0.57	0.17	0.25	0.79	−0.29	0.22	0.94	2.14	7.53	s2
	(2.54)	(3.87)	(3.83)	(6.23)	(−1.71)	(1.91)	0.83	(−1.14)	0.95	1967
Germany	0.68	0.11	0.27	0.20	0.49	−0.05	0.69	1.85	2.79	s2
	(2.04)	(1.49)	(4.02)	(1.33)	(3.74)	(−0.36)	0.57	(1.32)	3.02	1969
France	0.72	0.07	0.54	0.14	0.26	0.39	0.76	2.12	4.74	s2
	(1.89)	(1.27)	(4.63)	(0.99)	(1.93)	(2.80)	0.77	(1.55)	14.58	1970
Italy	0.83	0.23	0.30	0.51	0.08	0.25	1.29	2.04	2.84	s2
	(2.06)	(4.88)	(2.72)	(3.84)	(0.54)	(1.96)	0.81	(−0.93)	0.43	1964
United Kingdom	0.76	0.08	0.28	0.93	−0.27	0.17	1.31	2.07	3.01	s2
	(2.14)	(2.55)	(2.70)	(7.04)	(−1.50)	(1.35)	0.76	(−1.69)	3.21	1963
Canada[3]	1.20	0.14	0.26	0.61	0.00	0.02	0.88	1.66	7.02	s2
	(3.20)	(2.08)	(3.14)	(3.76)	(0.02)	(0.14)	0.69		3.25	1968
	1.19	0.07	0.14	0.60	0.00	−0.08	0.87	1.87	7.00	
	(3.32)	(1.24)	(3.49)	(3.81)	(0.03)	(−0.59)	0.70		3.23	
OECD total	0.64	0.44	0.81	0.61	0.25		0.50	1.66	6.61	1973
	(1.70)	(5.78)	(7.72)	(5.46)	(2.04)		0.96	(0.69)	7.98	
EMS	1.03	0.17	0.64	0.69	0.12		0.58	2.65	4.04	1971
	(1.79)	(1.66)	(4.97)	(3.89)	(0.64)		0.91	(2.58)	6.29	
Australia	2.47	0.12	0.65	0.91	−0.20		1.48	1.79	7.41	1963
	(3.72)	(0.94)	(3.09)	(4.99)	(−1.18)		0.84	(1.40)	3.86	
Austria[3]	2.15	0.15	0.42	0.57	0.02		1.20	1.69	1.82	1969
	(1.76)	(1.10)	(1.96)	(2.17)	(0.08)		0.57		1.30	
	2.70	0.12	0.23	0.38	−0.04		1.06	1.94	1.68	
	(2.13)	(1.45)	(2.99)	(1.56)	(−0.17)		0.67		5.17	
Belgium-Luxembourg	3.08	0.06	0.78	0.68	−0.22		1.49	1.89	6.60	1971
	(3.17)	(0.41)	(3.16)	(2.93)	(−1.00)		0.68		5.05	

Table 3. *(Continued)* **Regression results: unrestricted P* model**

Dependent variable: first difference of the log of the GDP deflator (Aln P)

	Intercept x 100	Velocity gap (–1)	output gap (–1)	Aln P (–1)	Aln P (–2)	SEE x 100 adj. R^2	DW (h-stat.)	Homog. F-test[1]	Start period[2]
Denmark	0.95	-0.14	0.57	0.69	0.17	1.31	1.67	0.55	1964
	(0.86)	(-2.10)	(3.98)	(3.87)	(0.97)	0.69	(1.96)	17.99	
Finland	1.73	0.73	0.65	0.81	-0.07	2.14	2.16	0.80	1963
	(1.19)	(3.40)	(3.42)	(4.91)	(-0.05)	0.66	(-1.71)	0.06	
Greece	2.12	0.44	1.30	0.43	0.46	2.14	2.03	3.28	1963
	(2.16)	(3.10)	(5.68)	(2.81)	(3.37)	0.90	(-0.19)	10.60	
Ireland[3]	3.57	0.36	0.59	0.45	0.18	3.46	1.92	2.83	1963
	(2.16)	(2.12)	(1.72)	(2.53)	(1.03)	0.53	(0.15)	0.23	
	8.56	0.17	0.64	0.08	-0.11	2.82	2.14	5.87	
	(4.68)	(1.30)	(3.61)	(0.47)	(-0.73)	0.69	(-1.11)	0.23	
Netherlands	2.08	0.14	0.11	0.67	-0.05	1.78	1.52	2.07	1963
	(1.93)	(2.01)	(1.77)	(3.43)	(-0.25)	0.63	(0.68)	2.65	
New Zealand	5.20	–0.09	0.15	0.74	-0.23	3.75	2.12	4.19	1967
	(2.63)	(-0.40)	(0.69)	(3.42)	(-1.07)	0.33		0.79	
Norway	1.93	0.23	0.66	0.47	0.23	3.16	2.15	0.56	1971
	(0.75)	(0.93)	(1.26)	(1.97)	(0.95)	0.13	(0.32)	0.41	
Spain3	4.21	0.25	0.64	0.51	0.14	2.05	1.84	4.11	1970
	(2.62)	(1.90)	(2.43)	(2.22)	(0.68)	0.77	(0.82)	1.77	
	3.38	0.22	0.20	0.38	0.26	1.81	2.22	6.84	
	(2.87)	(2.33)	(3.62)	(1.83)	(1.40)	0.83	(-1.53)	0.89	
Sweden	1.63	0.19	1.42	0.43	0.37	1.81	2.67	1.61	1963
	(1.71)	(1.77)	(5.31)	(3.03)	(2.57)	0.67	(-2.58)	21.37	
Switzerland	3.65	0.08	0.62	0.22	-0.16	1.00	1.56	7.79	1973
	(4.79)	(1.17)	(4.97)	(1.28)	(-1.15)	0.78	(1.17)	5.09	

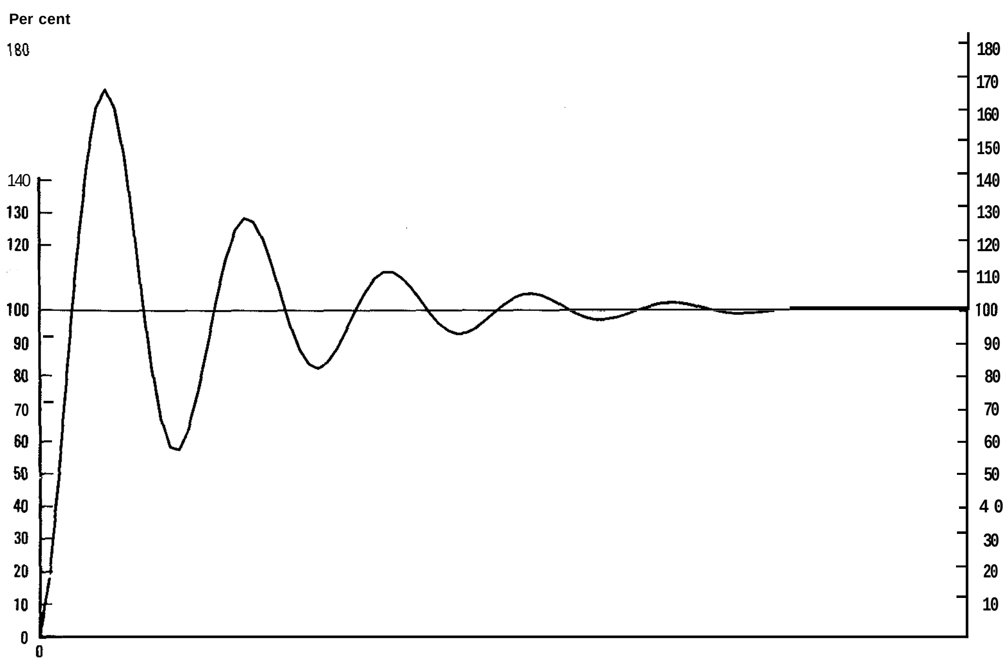
1. First line: F–test for the restriction that the intercept is equal to zero and the sum of price term coefficients is equal to unity.
   Second line: F–test for the restriction that the coefficient of the velocity gap and the output gap are equal.
2. End period = 1989.
3. In the second equation, trend, output and velocity are calculated using a linear time trend.

Among the countries for which both the velocity and the output gap matter, the data accepted the constraint of equality of the two gap coefficients only for Japan, Italy, Finland and Ireland. The other constraint of homogeneity of prices with respect to P* ($a_2 = 0$ and $a_3 + a_4 + a_5 = 1$) was accepted for the United States, Austria, Finland and Sweden[17]. The latter result should not be regarded as evidence against the P* model, however, as the change in the inflation rate was not zero on average over the estimation sample periods, whereas the velocity gap and output gap have a zero mean by construction.

Among the countries for which the velocity and output gaps were not significant are major energy (the Netherlands and Norway) and raw material producers (Australia and New Zealand). For these countries, movements in the GDP deflators are likely to be strongly influenced by changes in world market prices for commodities. New Zealand is, in addition, a country where temporary wage/price freezes were in place up to the mid-1980s. However, when dummy variables were added to the equations for these countries for the periods of the oil price shocks and wage/price freezes, the velocity gap coefficient did not become more significant.

Leaving aside the possibilities of the money supply being endogenous and reverse causation for countries with fixed exchange rates, shifts in monetary policy regimes may also be a factor behind equation instability. Several studies have found instability in wage and price equations in the 1980s (see Giavazzi and Giovannini, 1988, and

Chart 4. Adjustment of the price level to P-star
Cumulative response (%) to *a* level change in P-star



Poret, 1990, for France; Kremers, 1990, for Ireland; Weber, 1989, for Denmark and Belgium; Barell, 1990, for Italy). However, using interactive dummies did not provide evidence of a greater influence of the velocity gap since 1980 and 1983, respectively.

The best results, in terms of significance of the two gap coefficients and standard errors, were found for the OECD aggregate. In particular, the t-statistic and the coefficient value of the velocity gap are higher for the aggregate equation than the weighted average of single-country results. This could reflect greater stability of OECD-wide money demand as large portfolio shifts between countries tend to offset each other in the aggregate. This result is in line with the McKinnon currency substitution hypothesis (McKinnon, 1982; Viren, 1990).

Finally, the adjustment path of prices to P* is shown for the OECD aggregate equation in Chart **4.** While the response of inflation to a permanent shift in money supply is rapid, the price level overshoots initially after three half-years and adjusts to P* in damped oscillations. Full adjustment of the price level takes more than 31 years. Such oscillations are also found in large-scale econometric models (Hallman et *al.,* 1989). The equation suggests that steady inflation could be reached faster if the authorities were able to manipulate the money stock to offset price-level fluctuations after an initial shock. Oscillations of the price level around the equilibrium level would in any case argue against following a simple monetary rule in the presence of exogenous shocks.

IV. COMPARISON OF TRACKING AND FORECASTING PERFORMANCE OF ALTERNATIVE MODELS

In order to assess the tracking performance of the P* model against other inflation models, non-nested J-tests were undertaken for the countries where the estimated coefficient of the price gap is significant. The J-test examines whether there is information in a rival approach not contained in the model under investigation by adding the predicted values of the rival equation as regressors. If the estimated coefficients are significant, the model is said not to reject the rival equation.

As the novelty of the P* model is the addition of a velocity gap variable to the standard output gap model of inflation, the unrestricted P* model (based on the Hodrick-Prescott filter) has been tested against an output gap model, augmented by either an interest-rate term structure model or the T-bill model[18]. Hence, the test provides a comparison of the predictive power of alternative financial variables. Starting from equation [9], the velocity gap is replaced by the lagged yield gap $(R_{t-1} - r_{t-1})$ and the change in the short-term interest rate $(r_{t-1} - r_{t-2})$ in the case of the so-called T-bill model, where R and r refer to long and short-term interest rates, respectively. For most countries, R_t is the rate of return on a ten-year government bond and r_t the rate of return on three-month paper. The results suggest that the P* model is capable of rejecting the rival equations in eleven cases, whereas the converse is true in four cases only (Table **4).**

The forecasting accuracy of the P* approach against both rival models and published forecasts has been extensively researched in the United States. Results suggest

Table 4. In-sample non-nested tests[1]

	The unrestricted P* model is not rejected by its rival:		The P* model does not reject:	
	term-structure model	T-bill model	term-structure model	T-bill model
United States	3.0	2.3	1.9	-0.2
Japan	3.9	4.2	0.4	1.5
Germany	1.4	1.6	0.1	0.9
France	2.0	1.4	1.3	1.3
Italy	4.4	4.4	0.1	2.2
United Kingdom	1.8	1.8	0.7	0.0
Canada	2.8	1.9	2.5	0.2
Austria	0.5	0.5	0.5	0.8
Finland	1.4	1.7	-0.3	1.8
Greece	2.8	2.1	0.7	4.4
Ireland	2.6	1.8	2.7	0.7
Netherlands	1.2	1.5	1.5	0.5
Spain	2.5	2.9	0.8	1.1
Sweden	1.3	1.0	2.1	1.8
Switzerland	–0.9	0.5	2.5	1.5

1. The null hypothesis that model i is not rejected by model j is tested by adding the predicted values p_i of model i as regressors in model j. If the coefficient of p_i is significantly different from zero, model i is not rejected by model j. The t-statistics attached to this coefficient are reported in the table. (Davidson, R. and J.G. McKinnon (1981), "Several tests for model specification in the presence of alternative hypothesis", *Econometrica*, 49.) See text for details of rival model specification.

that the forecasts of the price-gap equation are superior to the alternatives (Christiano, **1989** and Hallman *et al.,* **1990).** The margin of superiority, however, is small against many rival forecasts.

Forecasting accuracy of the unrestricted P* model based on linear time trends has also been tested against output gap models where expectations are incorporated either by the yield gap or the T-bill model (as for the J-test above). Forecasts are of a rolling-horizon nature and the data are real time: the equation estimation period is extended successively from **1979** to **1988** in order to generate one-year-ahead forecasts from **1980** to **1989.** Tests were performed to establish whether the forecast errors of one model could be explained ("encompassed") by the forecasts of another model. The results suggest that the P* model encompasses its two rivals for Japan and Germany, while the opposite **is** true for Italy. The P* model also outperforms the term-structure model for Canada and Greece. For the other countries, no model seems to be superior to another in terms of forecasting performance (Table 5).

A comparison of forecasts of the P* model is also shown with a simple second-order auto regressive model and the OECD's one-year-ahead projections for **1982** to **1989,** as published in various issues of the *OECD Economic* ***Outlook*** and compiled by Ballis **(1989).** The price-gap model has a somewhat smaller mean absolute error than the official projection and the **AR(2)** model only for the United States and Germany

Table 5. Encompassing tests for real-time forecasting **performance**[1]

	P* (i) against term-structure model (j)	Term-structure model (i) against P* (j)	P* (i) against T-bill model (j)	T-bill model (i) against P* (j)
United States	**0.1**	**-1.0**	**-0.3**	**-1.0**
Japan	**-1.4**	**-4.9**	-1.1	**-3.3**
Germany	**0.3**	**-3.4**	**-1.4**	**-2.2**
France	**-0.7**	**-0.2**	**-0.7**	**-1.0**
Italy	**-3.9**	**-0.2**	**-2.6**	**-0.6**
United Kingdom	**-1.3**	**0.1**	**-0.9**	**-1.1**
Canada	**-1.8**	**-5.0**	**-1.3**	**-1.7**
Australia	**-4.3**	**2.3**	**-4.4**	**2.5**
Austria	**2.1**	**-2.8**	**0.0**	**-0.7**
Finland	**-0.8**	**-0.7**	**0.9**	**-1.9**
Greece	**-0.2**	**-2.1**	**0.4**	**-2.2**
Ireland	**-1.9**	**-0.8**	**-0.6**	**-1.1**
Netherlands	**-0.1**	**-0.7**	**-1.3**	**0.7**
Switzerland	-1.1	**-1.6**	**-5.2**	**-2.6**

1. The table shows t-statistics for β_{ij} in the regression:

$$\epsilon^i_t = \beta_{ij}(f^j_t - f^i_t) + u_t$$

where ϵ^i are out-of-sample residuals from model i, and f^i and f^j are out-of-sample predictions from models i and j, respectively. Model i is said to forecast encompass model j if β_{ij} is not significantly different from zero but β_{ji} is. (See Chong, Yock Y. and David F. Hendry (1986), "Econometric evaluation of linear macro-economic models," Review of Economic Studies, 53, pp. 671-90 (August).)

The out-of-sample period is from 1980 to 1989. The time-trend based unrestricted price-gap model is tested against a term-structure augmented-output-gap model, and a short-term interest rate augmented-output-gap model (called T-bill model). (See text for details).

(Table ***6).*** For the other countries, its forecast performance is worse and the differences are large in the cases of Japan, Italy and the United Kingdom when the Hodrick-Prescott filter is used. Choosing a simple output-gap model, with potential output proxied by a linear time trend, (column [5]) would have been equally good or better for forecasting inflation than any of the price-gap models shown in columns [1] to [4], except for Germany and Italy. Not surprisingly, using information from the full sample period for calculating trends, the P* model forecasts are better (last column).

While the coefficients of the rolling regressions change little between **1982** and **1989,** the value of P* during the forecast period is very different for some of the countries as compared with the P* calculated for the whole sample period. Trend changes in velocity and output were large in some Member countries and the Hodrick-Prescott filter simply mimics cyclical changes, if **no** future values are supplied (see such "backward" trends for velocity in Chart **3)**[19]. Thus, the difference between in and out-of-sample performance is mainly accounted for by the difficulty of discriminating ***ex*** ante between transitory changes and innovations in trend velocity and potential output.

Table 6. Rolling horizon forecasts for inflation

Mean absolute one-year ahead errors, 1982 – 1989, pour cent

	Price gap models				Output gap model		AR [2] model	OECD Economic Outlook	Price gap model (full sample estimate [1])
	[1]	[2]	[3]	[4]	[5]	[6]			
United States	0.7	1.6	1.5	0.7	0.7	1.4	1.0	0.9	0.5
Japan	2.0	3.3	3.8	1.6	1.0	2.5	1.9	1.0	0.6
Germany	0.8	0.8	0.8	0.8	1.2	1.0	1.0	0.7	0.6
France	1.4	1.6	1.4	1.4	0.9	1.6	1.4	0.6	0.6
Italy	1.7	3.0	**1.7**	1.7	2.9	1.3	1.3	1.2	0.6
United Kingdom	3.1	4.2	3.8	3.3	1.3	3.5	1.6	**1.2**	0.8
Canada	1.6	1.3	1.5	2.2	1.2	1.6	1.5	1.0	0.7
OECD	1.6	1.7	1.8	1.4	1.0	1.4	1.3	0.5	**0.3**

[1] Trend velocity and potential output calculated using linear time trends.
[2] Trend velocity and potential output calculated using the Hodrick-Prescott filter.
[3] Trend velocity calculated using a linear time trend and potential output using the Hodrick-Prescott filter.
[4] Trend velocity calculated using the Hodrick-Prescott filter and potential output using a linear time trend.
[5] Potential output calculated using a linear time trend.
[6] Potential output calculated using the Hodrick-Prescott filter.

V. CONCLUSIONS

The P* approach explains inflation in terms of tightness in both goods and financial markets. It states that in the long run the price level is determined by the money supply following the classic quantity-theoretic tradition. Goods and financial markets tightness are proxied by measures of output and velocity gaps. In this paper, trend values for output and velocity have been computed by several methods ranging from linear time trends to more sophisticated statistical filters. Tests of in-sample tracking ability and forecasting performance yielded the following conclusions:

i) For nine of the twenty OECD countries examined in this study and for the OECD in aggregate, P* equations track past inflation rates better than a model based solely on the output gap. Depending on one's viewpoint, this makes P-star's glass either half-full or half-empty. For most countries, however, the P* equations outperform other simple financial-market-based inflation models.

ii) The P* equations are less satisfactory for the purpose of short-term forecasting of inflation. Except for the United States and Germany, the P* forecasts are clearly dominated by predictions from an auto-regressive model or by OECD's official inflation forecasts. The poor forecasting performance of the P* approach is mainly due to the difficulty in discriminating ***ex ante*** between transitory and permanent changes in both trend velocity and potential output. Using the equations for forecasting purposes needs to be accompanied by judgement about both the tightness of product markets and current trends in velocity.

NOTES

1. The P* concept has subsequently been applied to analyse inflation pressures in Japan (Bank of Japan, 1990) and the United Kingdom (Hannah and James, 1989).
2. There is little empirical work on the nexus between monetary policy and potential output. Some studies suggest that fiscal and monetary policy can have lasting effects on output due to hysteresis effects (De Long and Summers, 1989), co-ordination failure (Durlauf, 1989) or imperfect competition and menu costs (Blanchard, 1990).
3. The P* approach can also be thought of as the reduced-form of inflation models embedded in large-scale econometric models. A boost to money growth in these models is likely to depress velocity below its trend and move actual output above potential via interest-rate and exchange-rate movements. The actual price level converges towards its new equilibrium level with a lag. The length of the lag depends on the interaction between financial and goods and labour markets, the degree of wage indexation, the lagged effects of induced exchange-rate change, etc. While these latter variables mainly affect the output gap, these models nowadays incorporate expectational effects in labour and asset market-price equations, so that there is a role for a direct and immediate effect of money growth during the adjustment process.
4. A new M3 series for France exists since 1978. It has been spliced with the old M2R series before 1978. Hence, regression results for France reported below need to be interpreted with caution.
5. However, there is an ongoing – and as yet unresolved – debate about the frequency (Perron, 1989) and the contribution of stochastic shocks to the overall variance of output (Cochrane, 1988; Cogley, 1990), and about the possibility of long-memory stationary processes (Campbell and Mankiw, 1989; Diebold and Rudebusch, 1989).
6. For instance, Boughton (1990) finds stable long-run money demand functions for large industrialised countries provided the restriction of long-run homogeneity of money with respect to prices is relaxed. Also Yoshida and Rasche (1990) find a stable combination **of** real M2+CD and real GNP at least until 1985 for Japan with an income-elasticity of money significantly greater than unity, leading them to reject the hypothesis that velocity is stationary.
7. Tatom (1990) also found a non-stationary velocity for M2 for the United States as well as a non-stationary price gap.
8. Mean reversion of aggregate OECD velocity is consistent with the finding of a stable aggregate money demand function for the EMS area as a whole but not for the individual EMS countries (Kremers and Lane, 1990). Aggregate stability may be due to the fact that frequent large cross-border portfolio shifts destabilise national relationships, while these shifts balance across countries.
9. A series x is said to be integrated of order d (written I(d)) if the dth difference of x is a stationary series. For example, a random walk, given by $x(t) = x(t-1) + e(t)$, where $e(t)$ is stationary, is I(1). A stationary (i.e. I(0)) series:

i) has finite variance which does not depend on time;

ii) is only temporarily affected by random innovations; and

iii) tends to fluctuate around the mean (which may include a deterministic trend).

10. The Hodrick-Prescott filter is described extensively in King and Rebelo **(1989)**. Technically, the trend, as calculated by the Hodrick-Prescott filter, minimises the sum of the squared deviations between a time-varying trend and the raw series under the constraint that the sum of the squared second differences (i.e. the acceleration of the trend) does not exceed a certain factor chosen by the user. In extracting the low-frequency component from the series, this filter uses both backward and forward observations. In this paper, in order to get reliable figures for the late 1980s, the money stock and output series were extended using the June 1990 ***OECD Economic Outlook*** projections.

11. The aggregate measure includes 20 countries accounting for **95** per cent of total OECD output. Inflation rates and capacity utilisation rates are aggregated using **1987** output weights, while average velocity is computed by multiplying national monetary aggregates and nominal GDP by **1987** exchange rates.

12. While lagged dependent variables were included among the regressors, the contemporaneous and lagged rates of change of P* were found to be insignificant, except for Canada, Australia, New Zealand and Spain for which an influence of the contemporaneous growth rate of P* was detected. For the sake of comparability with the results for the other countries, this term was not included in the equations for these four countries.

13. Homogeneity might still exist, even though the intercept is positive and the sum of the inflation coefficients is below unity, if inflation has no unit root (i.e. it tends to revert to a historical average, reflected in the intercept).

14. The output and velocity-gap terms when based on the Hodrick-Prescott filter are stationary variables by construction, while the inflation series may be integrated of order one. As a consequence, the t-statistics associated with the gap variables may not follow the usual t-distribution in an equation including **I(1)** variables. Using the homogeneity-constrained equations (that is, in terms of acceleration of inflation and without the intercept) did not change, however, the results for the countries for which the hypothesis of a unit root in inflation series was not rejected by the data.

15. As some studies suggest downward rigidity of prices, tests for an asymmetric influence of the velocity and output gaps depending on their sign have also been performed. They proved to be unsuccessful, however, and results are not reported.

16. The results for the United States are broadly in line with the Federal Reserve's study, except that the restriction of equality of the coefficients of the velocity and output gap is not accepted in our **U.S.** equation. However, using a similar output gap concept as Hallman ***et** al.* gives virtually the same results as they found.

17. The results for the United States are consistent with those of the Federal Reserve study, which found that the adjustment of the actual to the potential price level can be modelled by an inflation-acceleration equation.

18. Christiano **(1989)** investigated the out-of-sample performance of the P*, yield gap and T-bill models. For the latter two models, no output gap term was included. Blundell-Wignall ***et al.*** **(1990)** investigated the term structure as a leading indicator of inflation, also excluding an output gap term.

19. However, the use of the Hodrick-Prescott filter for calculating the velocity gap causes no more trouble in a forecasting context than the use of a deterministic time trend: combined with a time trend for potential output (column [4]), it yields better results for Japan and the OECD as a whole than the pure time trend-based price gap model (column [1]).

BIBLIOGRAPHY

Ballis, B. (1989), "A post mortem of OECD short-term projections from 1982 to 1987", *OECD Department of Economics and Statistics Working Papers,* No. 65 (February).

Bank of Japan, Research and Statistics Department (1990), "A study of potential pressure on prices. Application of P* to the Japanese economy", Special Paper No. 186 (February).

Barell, R. (1990), "Has the EMS changed wage and price behaviour in Europe?", *National Institute Economic Review* (November).

Blanchard, O.J. (1990), "Why does money affect output? A survey", in B.M. Friedman and F.H. Hahn (eds.), *Handbook of Monetary Economics,* Vol. II, Elsevier Science Publishers B.V.

Blundell-Wignall, A., F. Browne and P. Manasse (1990), "Monetary policy in the wake of financial liberalisation", *OECD Economic Studies,* No. 15 (Autumn).

Bomhoff, E.J. (1990), "Stability of velocity in the group of seven countries: a Kalman filter approach", *IMF Working Paper* (WP/90/80) (September).

Boughton, J.M. (1990), "Long-run money demand in large industrial countries", *IMF Working Paper* (WP/90/53) (June).

Braun, S. (1990), "Estimation of current quarter gross national product pooling preliminary labor market data," *Journal of Business and Economic Statistics,* 8, 293-304 (July).

Browne, F.X. (1986), "A monthly money market model for Ireland in the EMS", *Central Bank of Ireland,* Annual Report (Spring).

Christiano, L.J. (1989), "P*: Not the inflation forecaster's holy grail", *Federal Reserve Bank of Minneapolis, Quarterly Review* (Fall).

Campbell, J.Y. and N.G. Mankiw (1989), "International evidence on the persistence of economic fluctuations", *Journal of Monetary Economics,* 23.

Cogley, T. (1990), "International evidence on the size of the random walk in output", *Journal of Political Economy,* Vol. 98, No. 3.

Cochrane, J.H. (1988), "How big is the random walk in GNP?", *Journal of Political Economy,* Vol. 96, No. 5.

De Long, J.B. and L.H. Summers (1989), "How does macroeconomic policy affect output?", *Brookings Papers on Economic Activity,* 2.

Diebold, F.X. and G.D. Rudebusch (1989), "Long memory and persistence in aggregate output", *Journal of Monetary Economics,* 24.

Durlauf, S.N. (1989), "Output persistence, economic structure, and the choice of stabilisation policy", *Brookings Papers on Economic Activity,* 1.

Ebrill, L.P. and S.M. Fries (1990), "The dynamics of money demand and prices", *IMF Working Paper* 90/75 (August).

Giavazzi, F. and A. Giovannini **(1988),** "The role of the exchange-rate regime in a disinflation: empirical evidence on the European Monetary System" in Giavazzi, F., S. Micossi and M. Miller (eds.), *The European Monetary System,* Cambridge University Press.

Gordon, R.J. **(1990),** "What is New-Keynesian economics?" *Journal of Economic Literature,* No. **3** (September).

Hallman, J.J., R.D. Porter and D.H. Small **(1990), "Is** the price level tied to the stock of M2 in the long run?", Board of Governors of the Federal Reserve System, mimeo.

Hallman, J.J., R.D. Porter and D.H. Small **(1989),** "M2 per unit of potential GNP as an anchor for the price level", *Staff Study,* No. **157,** Board of Governors of the Federal Reserve System (April).

Hannah, S. and A. James **(1989),** "P-star as a monetary indicator for the UK", *NatWest Capital Markets* (June).

Haslag, J.H. **(1990),** "Monetary aggregates and the rate of inflation", *Federal Reserve Bank of Dallas Economic Review* (March).

Humphrey, T.M. **(1989),** "Precursors of the P-star model", *Federal Reserve Bank of Richmond Economic Review* (July/August).

King, R.G. and S.T. Rebelo **(1989),** "Low frequency filtering and real business cycles", *Discussion Papers,* Rochester Centre for Economic Research, No. **205** (October).

Kremers, J.M. **(1990),** "Gaining a policy credibility for a disinflation", *IMF Staff Papers* (March).

Kremers, J.M., and T. Lane **(1990),** "Economic and monetary integration and the aggregate demand for money, in the EMS', *IMF Working Paper,* No. **WP/90/23** (March).

Kuttner, K.N. **(1990),** "Inflation and the growth rate of money", *Economic Perspectives,* Federal Reserve Bank of Chicago (January/February).

McKinnon, R.I. **(1982),** "Currency substitution and instability in the World Dollar standard", *American Economic Review,* No. **72.**

Pecchenino, R.A. and R.H. Rasche **(1990),** "P* type models: evaluation and forecasts", *NBER Working Papers Series,* No. **3406** (August).

Perron, P. **(1989),** "The great crash, the oil price shock and the unit root hypothesis," *Econometrica,* Vol. **57,** No. 6 (November).

Perron, P. **(1988),** "Trends and random walks in macroeconomic time series: further evidence from a new approach", *Journal* ***of*** *Economic Dynamics* and *Control,* No. 12.

Poret, P. **(1990),** "The puzzle of wage moderation in the **1980s",** *OECD Department of Economics and Statistics Working Papers,* No. **87** (December).

Tatom, J.A. **(1990),** "The P-star approach to the link between money and prices", Federal Reserve Bank of St. Louis, mimeo.

Torres, R. and J.P. Martin **(1990),** "Measuring potential output in the seven major OECD countries", *OECD Economic Studies,* No. **14** (Spring).

Viren, M. **(1990),** "McKinnon's currency substitution hypothesis: some new evidence", *Economic International,* Vol. **XLIII,** No. **2-3** (May/August).

Weber, A.A. **(1989),** "The role of policymakers reputation in the ems-deflation: an empirical **evaluation",** revised version of the Centre of Economic *Research Discussion Paper,* No. **8803,** Tilburg University.

Yoshida, T. and R.H. Rasche **(1990),** "The M2 demand in Japan: shifted and unstable?", *Bank of Japan Monetary and Economic Studies* (September).