OECD Economics Department Working Papers No. 101

P-Star as an Indicator of Inflationary Pressure

Peter Hoeller,
Pierre Poret

https://dx.doi.org/10.1787/821307064337

OECD

General Distribution
OCDE/GD(91)93

OECD
DEPARTMENT OF ECONOMICS AND STATISTICS

WORKING PAPERS

N°. 101 P-STAR AS AN INDICATOR OF INFLATIONARY PRESSURE

by

Peter Hoeller and Pierre Poret
Public Economics Division

March 1991

OCDE
OECD

GENERAL DISTRIBUTION

OCDE/GD(91)93

OECD

DEPARTMENT OF ECONOMICS AND STATISTICS

WORKING PAPERS

WORKING PAPER NO. 101

P STAR AS AN INDICATOR OF INFLATIONARY PRESSURE

by
Peter Hoeller and Pierre Poret
Public Economics Division

ORGANISATION FOR ECONOMIC CO-OPERATION AND DEVELOPMENT

Paris 1991

35067

GENERAL DISTRIBUTION

ECONOMICS AND STATISTICS DEPARTMENT

WORKING PAPERS

This series of Working Papers is designed to make available, to a wider readership, selected studies which the Department has prepared for use within OECD. Authorship is generally collective, but main individual authors are named. The Papers are generally available in their original language, English or French, with a summary in the other.

Comment on the Papers is invited, and may be sent to OECD, Department of Economics and Statistics, 2 rue André Pascal, 75775 Paris Cedex 16, France. Additional copies of the Papers on a limited basis can be forwarded on request.

ORGANISATION FOR ECONOMIC CO-OPERATION AND DEVELOPMENT

The P-star approach has been developed by the U.S. Federal Reserve as a new indicator of inflationary pressures. This paper assesses its usefulness for 20 OECD Member countries. Regression results are presented and in-sample tracking ability and forecasting performance of the equations are compared to rival inflation models and official OECD projections.

* * * * * *

L'approche dite P-star a été développée par la Réserve Fédérale des États-Unis comme un nouvel indicateur des pressions inflationnistes. Ce papier examine son utilité pour 20 pays de l'OCDE. Des résultats de regression sont présentés ; la capacité des équations à simuler et prévoir l'évolution des prix est comparée à celle de modèles rivaux et aux projections officielles d'inflation de l'OCDE.

N°. 101 P-STAR AS AN INDICATOR OF INFLATIONARY PRESSURE

by

Peter Hoeller and Pierre Poret*

Public Economics Division

March 1991

* The authors are grateful to Frank Browne, Andrew Dean, Jorgen Elmeskov, Andreas Fischer, Robert Ford, John Martin, Robert Price, John Roberts and Peter Sturm for comments on earlier versions of this paper. The authors are indebted to Laurence Le Fouler for technical assistance.

CONTENTS

TABLES

CHARTS

P-STAR AS AN INDICATOR OF INFLATIONARY PRESSURE

I. INTRODUCTION AND SUMMARY

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, past 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 approach for policy implementation, therefore, depends on the ability of the monetary authorities to influence monetary aggregates in the short and long run (Pecchenino and Rasche, 1990). Furthermore, for the P* concept to be truly useful, it must provide information not captured by other inflation models (Christiano, 1990; 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 II describes the theoretical framework. Section III 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. In-sample tracking ability and forecasting performance of the equations are then compared to rival inflation models and to official OECD projections.

The main results are the following:

i) Satisfactory price-gap equations are found for seventeen out of twenty OECD countries; allowing for stochastic trends yields a significant influence of the velocity gap component - the main novelty of the P* model - in ten countries;

ii) For the latter countries, statistical tests suggest that P* equations track historical data better than rival inflation models;

iii) However, rolling-horizon forecasts exploiting only past information show that forecast errors of the price-gap equations would have been much worse than the official OECD projections in the 1980s for most large OECD countries. The difference between in- and out-of-sample performance is mainly accounted for by the difficulty of discriminating <u>ex ante</u> between transitory changes and innovations in trend velocity and potential output.

II. 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:

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

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

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

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:

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

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^*. For this, one needs an hypothesis about the adjustment process. In the Federal Reserve study, the adjustment of the actual to the potential price level is modeled by an inflation-acceleration equation:

[4] $dp - dp_{-1} = a\ (p^*_{-1} - p_{-1})$

With $a > 0$, is the speed of adjustment of prices to P^*.

In equation [4], the two components of the price gap, the velocity and output gaps, are constrained to have the same coefficient. It is interesting to consider both gaps separately as in equation [5]:

[5] $dp - dp_{-1} = a_1(v^*_{-1} - v_{-1}) + a_2(q_{-1} - q^*_{-1}) \quad ; \quad a_1,\ a_2 > 0$

New-Keynesian approaches towards modelling inflation focus on the output gap and inertia in price adjustment (Gordon, 1990). In these models, inflation accelerates when the output gap opens up, while the price level is indeterminate ($a_1=0$, $a_2>0$). Such an inflation model is a standard expectations-augmented Phillips curve, where expectations (dp^e) are assumed to be adaptive and represented by past inflation:

[6] $dp = a_2(q_{-1} - q^*_{-1}) + dp^e$

The expectations-augmented Phillips curve and the price-gap model coincide if the set of information available to agents is assumed to include the velocity gap (3):

[7] $dp^e = dp_{-1} + a_1 (v^*_{-1} - v_{-1})$

The usefulness of P* 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, although prices may still lag the domestic money stock.

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.

III. 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 variable theoretically termed "money". The money aggregates projected during OECD's half-yearly forecasting exercises are used here. Thus, 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) linear time trends; and (ii) the low-frequency component 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 deterministic time trend over the sample periods. The null hypothesis of non-stationarity could be rejected only in the cases of Austria, Canada, Switzerland, Spain and the OECD in aggregate (7).

The likely 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 (8). 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 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 cointegrating regression of velocity on the fraction of the labour force employed in agriculture (as a proxy for the industrialisation and 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 more unpredictable in the 1980s for the large OECD countries. For this paper, we prefer to use the Hodrick-Prescott filter (9), 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 twenty countries are examined (10). Also the regression results reported below proved to be slightly better when the Hodrick-Prescott smoothed real output was used rather than a more sophisticated measure of potential derived from the supply block of the OECD's INTERLINK model (Torres and Martin, 1990).

The split of output into "transitory" and "permanent" 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.

Once potential output and trend velocity are defined, price gaps are calculated according to equation [3]. For illustrative purposes, the increase in 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 changes are between one and two years. Following a prolonged period of a negative price gap in the early to mid-1980s accompanied by falling inflation, the price gap became positive in 1987 and inflation has since drifted up again. In 1989 the aggregate price gap pointed to a stable future inflation development. 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 (Hodrick-Prescott filter based) are shown for the seven large countries in Charts 2 and 3.

IV. 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:

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

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

The dependent variable is the inflation rate. The first term on the right-hand side of equation [8] is the price gap; in equation [9] this is split up into its two components (12). If the intercept, a_0, is zero and the coefficients on lagged inflation sum to unity, identity between actual and potential price levels in the long-run will be achieved (13). If, in addition, a_1 and a_2 in equation [9] do not differ significantly, the equation could be parameterised more parsimoniously, using only the price gap as in equation [8]. In this case, 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 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, Norway and New Zealand, 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 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 Ireland, the Netherlands, Norway and New Zealand, the coefficient of the velocity gap was insignificant for Germany, France, Belgium, Denmark, Netherlands, Norway, Australia, Switzerland and New Zealand and wrongly signed in two of these countries (15). Out of the 20 countries examined, the velocity gap and output gap have correctly signed and significant coefficients for ten (16).

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, Ireland and Finland. The other constraint of homogeneity of prices with respect to P* ($a_0=0$ and $a_3+a_4+a_5=1$) was accepted for the United States, Austria, Finland and Sweden. 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 (Norway and the Netherlands) 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 often 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.

Apart from the problem of endogenity of the money supply and possible reverse causation for countries with fixed exchange rates, shifts in monetary policy regimes can also be a factor behind equation instability. A number of studies found instability in wage and price equations in the 1980s (see Giavazzi and Giovannini (1988) and 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 takes more than 31 years. Such oscillations are also found in large-scale econometric models (Hallman et al., 1989). Steady inflation can be reached faster if the money stock is manipulated 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.

V. 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 they 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 (17). 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 rival models and against published forecasts has been extensively researched in the United States. Results suggest that the forecasts of the price-gap equation are superior to the alternatives (Christiano, 1989 and Hallman <u>et al.</u>, 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. Similar to the J-test performed above, it is tested whether forecast errors of one model can 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 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 a linear time trend for proxying potential output (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 for both ways of calculating P^*. 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) (18). 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.

NOTES

1. The P* concept has also been applied to analyse inflation pressures in Japan (Bank of Japan, 1990) and the United Kingdom (Hannah and James, 1989).

2. This assumption is controversial. Recent 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), imperfect competition and menu costs (Blanchard, 1989).

3. The P* approach can also be thought of as the reduced-form of inflation acceleration mechanisms inherent 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 a 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 Rudebush, 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, so that they reject the hypothesis that velocity is stationary.

7. Mean reversion of aggregate 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.

8. Tatom (1990) found a non-stationary velocity for M2 for the United states as well as a non-stationary price gap. While the gap was found to be stationary in first difference, the M2 based, first-differenced, P^* measure turned out to be statistically insignificant in explaining the level of prices. On the other hand, Tatom found a significant link between M1 growth and price changes.

9. 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 raw series were extended using the June 1990 <u>OECD Economic Outlook</u> projections.

10. Lucas (1980) applied a two-sided exponentially-weighted moving-average filter, very similar to the Hodrick-Prescott filter, to the inflation rate and money growth to investigate empirical implications of the quantity theory of money. The rationale for his smoothing approach was that slow-moving structural changes are well understood by agents but cannot be observed by the econometrician. According to Lucas, the hope in applying such a filter is not that the underlying model holds exactly but that the actual data series are generated by a very slowly changing structure of the financial system, while business-cycle activity is occurring at higher frequencies superimposed. Similarly, the assumption implicit in the filtering approach here is that permanent shocks due to financial market deregulation and innovation have only gradual effects on long-run velocity, as the use of the Hodrick-Prescott filter does not allow for abrupt shifts in trend velocity.

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, Spain, Australia and New Zealand 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 non-stationary. As a consequence, the t-statistics associated with the gap variables may not follow the usual t-distribution. 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. This is probably due to insufficient number of observations. When using quarterly rather than semi-annual data, evidence for asymmetric inflation-output trade-offs can be found for some OECD countries (Poret, 1991).

16. The results for the United States are broadly in line with the Federal Reserve's study, except that the restriction of equality between 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. Christiano (1988) 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.

18. 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]).

REFERENCES

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 (ed.), 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 Mankiw, N. G., (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 Summers, L.H., (1988), "How does macroeconomic policy affect output?", Brookings Papers on Economic Activity, 2.

Diebold, F. X. and Rudebusch, G.D., (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 (1989), "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.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).

Lucas, R.E. (1980), "Two illustrations of the quantity theory of money", American Economic Review, Vol. 70, No. 5, (December).

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).

Poret, P. (1991), "Has disinflationary policy gained credibility? A quick investigation of the inflation-output trade-off in the 1980s", mimeo, OECD.

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 Rasche, R.H., (1990), "The M2 demand in Japan: Shifted and Unstable?", BOJ Monetary and Economic Studies, September.

Table 1

Unit root tests: logarithm of money velocity

			Alternative 3		Alternative 2	Alternative 1
	T	Start period	Φ3	Φ2	Φ1	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
New Zealand	(48)	S1 1966	3.01	2.25	0.40	-0.84
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
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.

<u>Description of the test</u>

The unit root tests reported in Table 1 are augmented Dickey-Fuller tests with second-order correction. The testing strategy, following Perron (1988), involves a sequence of tests that run from general to restricted alternative hypotheses. We begin with the alternative hypothesis of a stationary series with a time trend:

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

If $\beta = 0$, there is no time trend; if $\alpha = 0$ 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 Φ_3 jointly tests the two zero restrictions ($\alpha=\beta=0$) 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 Φ_2 therefore jointly tests three zero restrictions ($\mu=\alpha=\beta=0$). 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 Φ_3, 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 Φ_1 jointly tests two zero restrictions ($\mu=\alpha=0$). 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 ($\alpha=0$), 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 Φ_3 is below, <u>and</u> Φ_2 is above, its critical value, is the series deemed to have a unit root.

In Table 1, any test sequence that concludes that the series is stationary is marked with an asterisk. For 50 observations, the critical value for the t-statistic at the 5 per cent is -1.95 (Fuller (1976), <u>Introduction to Statistical Time Series</u>, p. 373). Critical values for Φ_3, Φ_2 and Φ_1 are 6.73, 5.13, 4.86 (Dickey and Fuller (1981), <u>Econometrica</u> 4, p. 1063).

Table 2

Regression results: restricted P* model

Dependent variable: first difference of the log of the GDP deflator ($\Delta \ln P$)

		Intercept x 100	Price gap (-1)	$\Delta \ln P$ (-1)	$\Delta \ln P$ (-2)	$\Delta \ln P$ (-3)	SEE x 100 adj. R^2	DW (h)	Start period (c)
United States	(a)	0.16 (0.83)	0.07 (3.56)	0.69 (5.50)	0.16 (1.03)	0.07 (0.54)	0.58 0.78	2.08 (-1.44)	S1 1962
	(b)	-0.00 (-0.03)	0.12 (3.97)	0.72 (5.83)	0.21 (1.33)	0.08 (0.62)	0.57 0.79	2.06 (-0.88)	
Japan	(a)	0.81 (3.36)	0.09 (5.36)	0.75 (5.45)	-0.32 (-1.84)	0.02 (0.15)	0.97 0.81	2.10 (-0.90)	S2 1967
	(b)	0.56 (2.51)	0.20 (5.81)	0.80 (6.33)	-0.27 (-1.62)	0.20 (1.77)	0.94 0.83	2.13 (-0.98)	
Germany	(a)	0.84 (2.60)	0.10 (2.90)	0.27 (1.83)	0.42 (2.98)	-0.21 (-1.43)	0.75 0.50	2.04 (-0.52)	S2 1969
	(b)	0.37 (1.22)	0.19 (3.45)	0.32 (2.23)	0.55 (4.15)	-0.08 (-0.59)	0.72 0.53	2.07 (-0.66)	
France	(a)	1.40 (2.80)	0.09 (3.65)	0.22 (1.43)	0.15 (1.01)	0.18 (1.28)	0.84 0.73	2.01 (-0.81)	S2 1970
	(b)	0.62 (1.40)	0.16 (2.94)	0.37 (2.44)	0.23 (1.48)	0.23 (1.49)	0.89 0.70	2.04 (-1.38)	
Italy	(a)	1.43 (3.21)	0.07 (4.39)	0.49 (3.71)	0.01 (0.09)	0.16 (1.34)	1.36 0.78	2.09 (-1.09)	S2 1964
	(b)	0.75 (2.0)	0.24 (5.40)	0.54 (4.50)	0.10 (0.72)	0.21 (1.84)	1.27 0.81	2.08 (-0.67)	
United Kingdom	(a)	0.66 (1.61)	0.03 (1.77)	1.07 (7.62)	-0.39 (-1.96)	0.19 (1.31)	1.45 0.71	2.03 (0.73)	S2 1963
	(b)	0.69 (1.89)	0.11 (3.40)	0.98 (7.40)	-0.30 (-1.60)	0.16 (1.21)	1.34 0.75	2.02 (-1.01)	
Canada	(a)	1.12 (3.31)	0.12 (3.30)	0.63 (3.99)	0.27 (0.14)	-0.11 (-0.79)	0.87 0.70	1.85 (0.73)	S2 1968
	(b)	1.18 (3.01)	0.17 (2.70)	0.67 (4.13)	0.04 (0.18)	-0.06 (-0.40)	0.90 0.68	1.75 (0.71)	
OECD total	(a)	2.52 (3.45)	0.24 (5.33)	0.73 (4.61)	-0.20 (-1.37)	.. ..	0.87 0.87	1.76 0.68	1973
	(b)	1.12 (2.01)	0.46 (6.17)	0.91 (6.97)	-0.11 (-0.77)	.. ..	0.78 0.89	1.69 (-0.18)	
EMS	(a)	2.92 (3.41)	0.16 (4.10)	0.83 (4.76)	-0.34 (-2.10)	.. ..	0.73 0.86	2.11 (-0.46)	1971
	(b)	1.06 (1.58)	0.36 (4.79)	1.01 (6.85)	-0.19 (-1.22)	.. ..	0.67 0.88	2.17 (-0.78)	

a) Price gap calculated using linear time trends.
b) Price gap calculated using the Hodrick-Prescott filter.
c) End period = 1989.

Table 2 (Continued)

Regression results: restricted P* model

Dependent variable: first difference of the log of the GDP deflator ($\Delta \ln P$)

		Intercept x 100	Price gap (-1)	$\Delta \ln P$ (-1)	$\Delta \ln P$ (-2)	SEE x 100 adj. R^2	DW (h)	Start period (c)
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)	
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)	
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)	
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)	
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)	
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.
c) End period = 1989.

Table 3

Regression results: unrestricted P^* model
Dependent variable: first difference of the log of the GDP deflator ($\Delta \ln P$)

	Intercept x 100	Velocity gap (-1)	Output gap (-1)	$\Delta \ln P$ (-1)	$\Delta \ln P$ (-2)	$\Delta \ln P$ (-3)	SEE x 100 adj. R^2	DW (h-stat.)	Homog. F-test (a)	Start period (b)
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*	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	

a) 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.

b) End period = 1989.

* In the second equation, trend and output velocity are calculated using a linear time trend.

Table 3 (Continued)

Regression results: unrestricted P^* model
Dependent variable: first difference of the log of the GDP deflator ($\Delta \ln P$)

	Intercept x 100	Velocity gap (-1)	Output gap (-1)	Δln P (-1)	Δln P (-2)	SEE x 100 adj. R^2	DW (h-stat.)	Homog. F-test (a)	Start period (b)
Austria*	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	
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*	3.57	**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	
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	
Spain*	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	
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	
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	
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	

a) 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.

b) End period = 1989.

* In the second equation, trend and output velocity are calculated using a linear time trend.

Table 4

In-sample non-nested tests (a)

	The unrestricted P^* model is not rejected by its rival: term-structure model	The unrestricted P^* model is not rejected by its rival: T-bill model	The term-structure model is not rejected by the P^* model	The T-bill model is not rejected by the P^* model
Jnited 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
Jnited 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

a) 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 McKinnon, J.G., "Several tests for model specification in the presence of alternative hypothesis", Econometrica, 49, 1981.) See text for details of rival model specification.

Table 5

Encompassing tests for real-time forecasting performance (a)

	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)
nited States	0.1	-1.0	-0.3	-1.0
apan	-1.4	-4.9	-1.1	-3.3
ermany	0.3	-3.4	-1.4	-2.2
rance	-0.7	-0.2	-0.7	-1.0
taly	-3.9	-0.2	-2.6	-0.6
nited Kingdom	-1.3	0.1	-0.9	-1.1
anada	-1.8	-5.0	-1.3	-1.7
ustria	2.1	-2.8	0.0	-0.7
inland	-0.8	-0.7	0.9	-1.9
reece	-0.2	-2.1	0.4	-2.2
reland	-1.9	-0.8	-0.6	-1.1
etherlands	-0.1	-0.7	-1.3	0.7
ustralia	-4.3	2.3	-4.4	2.5
witzerland	-1.1	-1.6	-5.2	-2.6

a) **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, "Econometric evaluation of linear macro-economic models," Review of Economic Studies, August 1986, 53, 671-90).

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

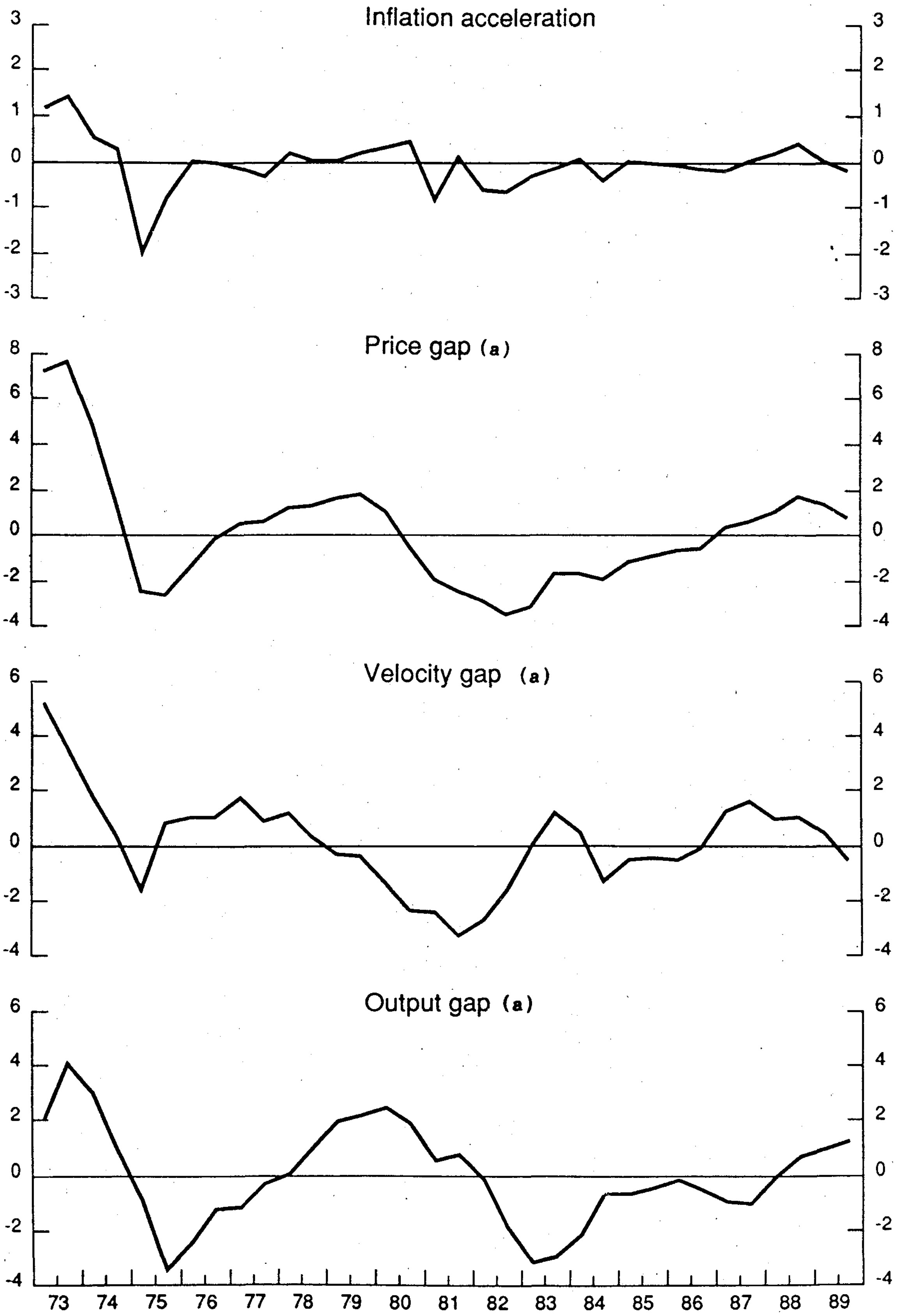
Rolling horizon forecasts for inflation
(Mean absolute one-year ahead errors, 1982 - 1989, per cent)

	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

Chart 1

The price gap and its components
for the OECD aggregate



a) Lagged by one period

Chart 2. Actual and trend output
(in logarithm)

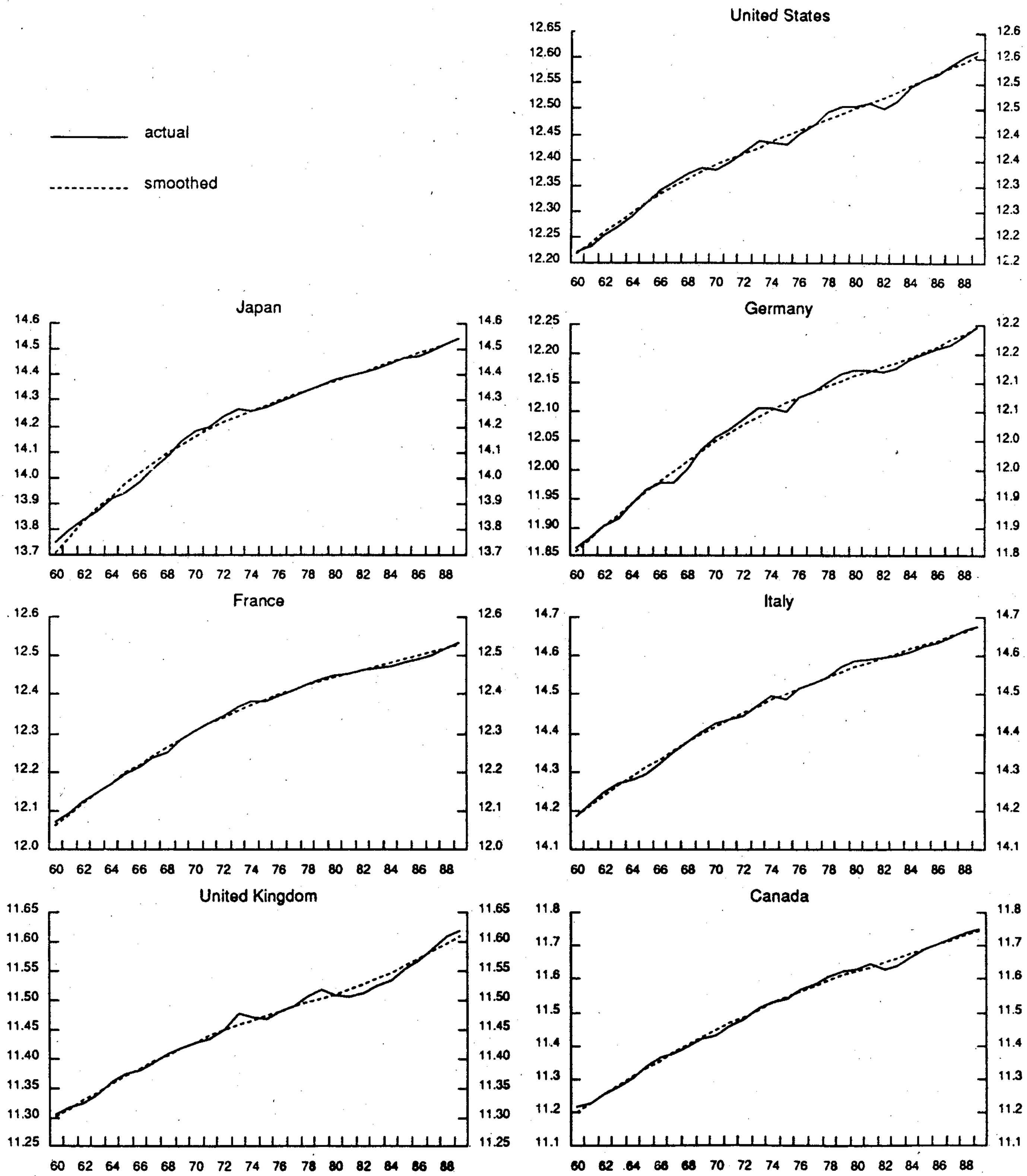
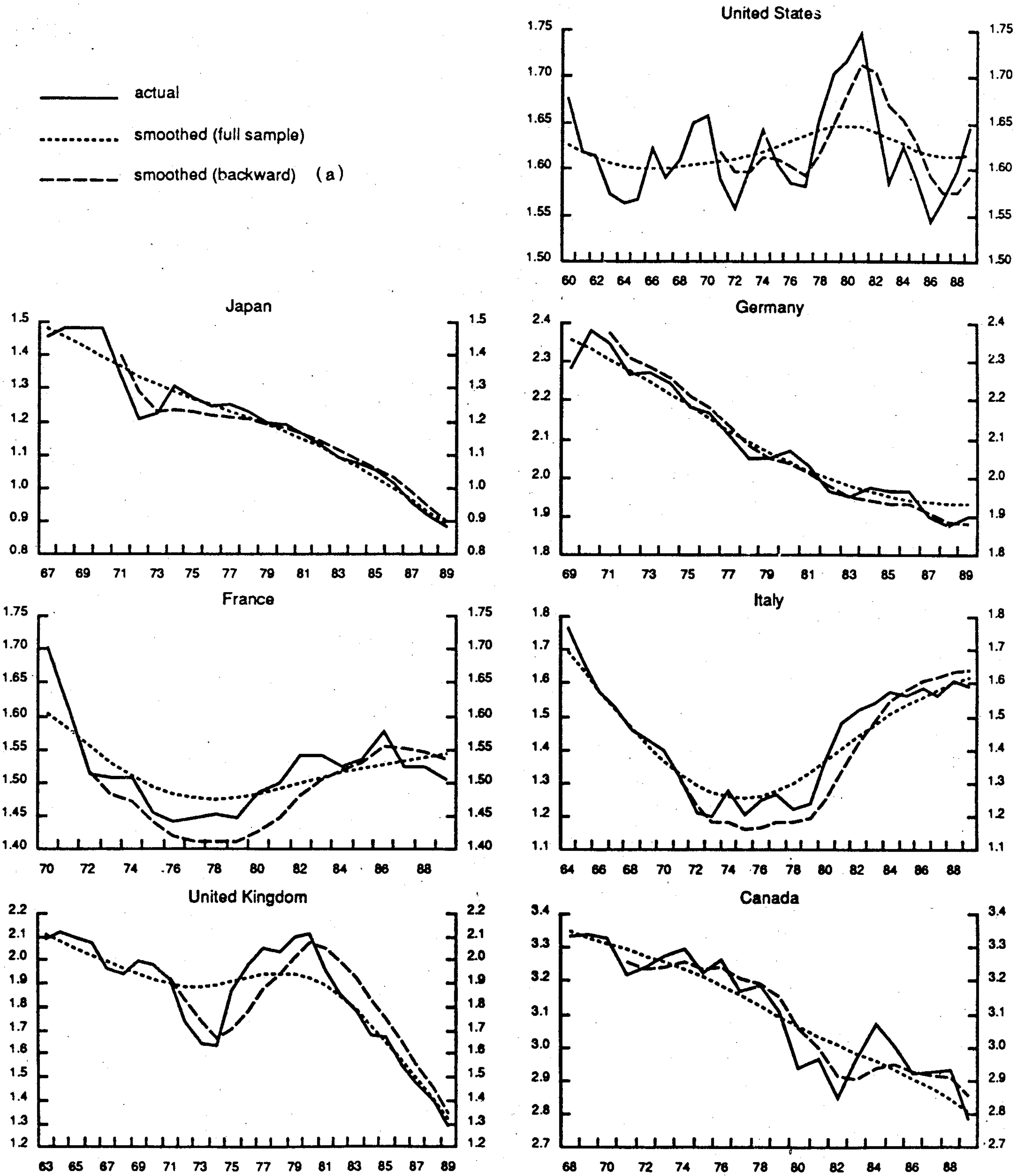


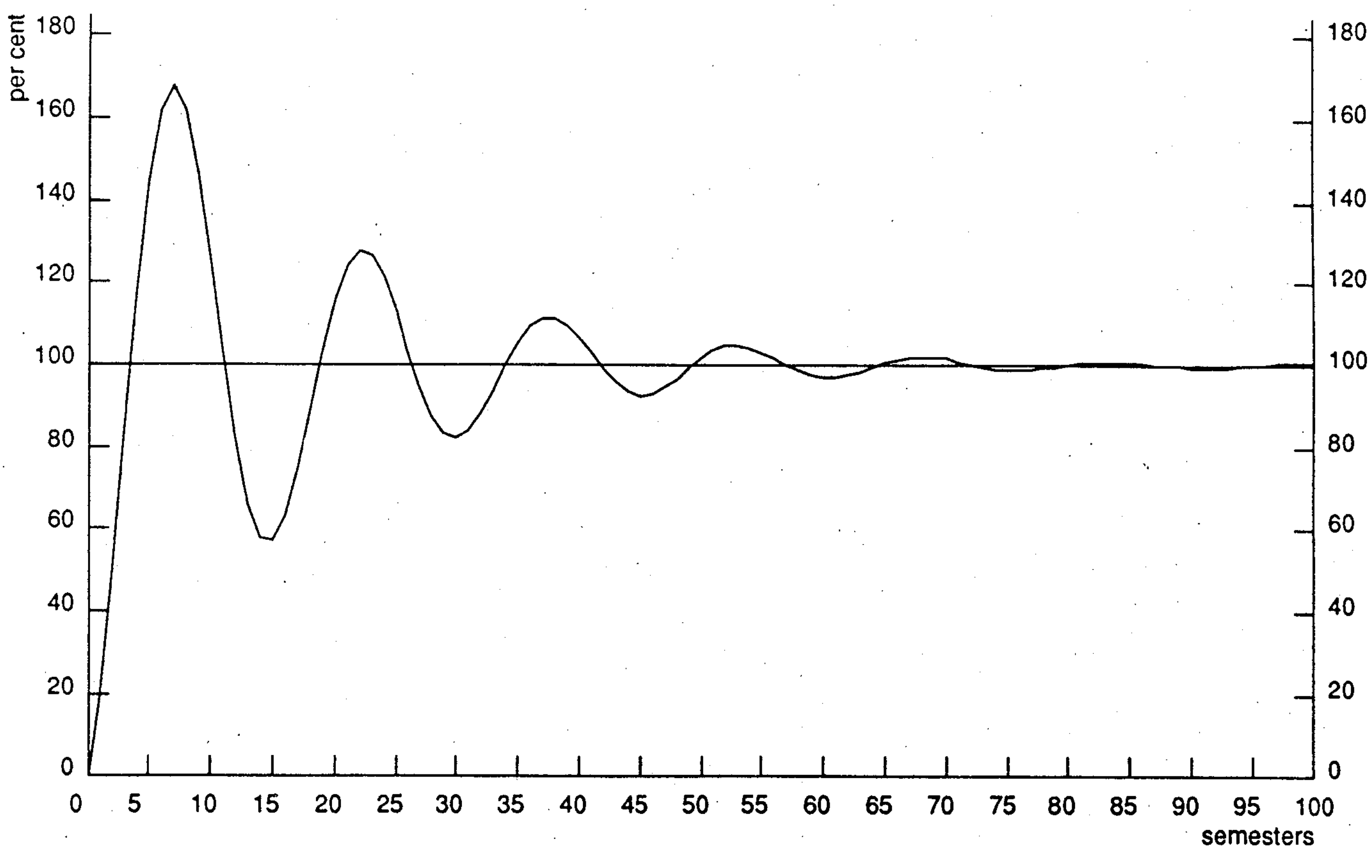
Chart 3. Actual and trend velocity



a) 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.

Chart 4. Adjustment of the price level to P star

(Cumulative response (%) to a level change in P-star)



ECONOMIC AND STATISTICS DEPARTMENT

WORKING PAPERS

A complete list of Working Papers is available on request.
A complete list of Working Papers No's 1-97 is available in Working Paper No. 98.

100. Old Age Income Maintenance
Murray Petrie, Peter Sturm

99. The Measurement of Output and Factors of Production for the Business Sector in the 24 OECD Member Countries (March 91)
Mark Keese, Gerard Salou, Pete Richardson

98. Macroeconomic Consequences of Financial Liberalisation: A Summary Report (February 1991)
Adrian Blundell-Wignall and Frank Browne

97. Deregulation, Credit Rationing, Financial Fragility and Economic Performance (February 1991)
Michael Driscoll

96. Increasing Financial Market Integration, Real Exchange Rates and Macroeconomic Adjustment (February 1991)
Adrian Blundell-Wignall and Frank Browne

95. Financial Liberalisation, Asset Prices and Exchange Rates (February 1991)
Marcus Miller and Paul Weller

94. Financial Liberalisation and International Trends in Stock, Corporate Bond and Foreign Exchange Market Volatilities (February, 1991)
Paul Kupiec

93. A macroeconomic model for Debt Analysis of the Latin America region and Debt Accounting models for the highly indebted countries (February, 1991)
Peter Dittus and Paul O'Brien

92. Unemployment Persistence and Insider-Outsider Forces in Wage Determination (February, 1991)
Bertil Holmlund

91. Infrastructure and Private-Sector Productivity (January 1991)
Robert Ford and Pierre Poret

90. The Public Sector: Issues for the 1990s(December, 1990)
Howard Oxley, Maria Maher, John P. Martin, Giuseppe Nicoletti and Patricia Alonso-Gamo

89. A Survey of Studies of the Costs of Reducing Greenhouse Gas Emissions (December 1990)
Peter Hoeller, Andrew Dean and Jon Nicolaison

88. Business investment in the OECD economies: recent performance and some implications for policy (November 1990)
Robert Ford, Pierre Poret

87. The "PUZZLE" of wage moderation in the 1980s (November 1990)
Pierre Poret

86. Modelling wages and prices for the smaller OECD countries (October 1990)
Kenichi Kawasaki, Peter Hoeller, Pierre Poret

85. Simulating the OECD INTERLINK Model under Alternative Monetary Policy Rules (October 1990)
Pete Richardson

84. WALRAS -- A multi-sector, multi-country applied general equilibrium model for quantifying the economy-wide effects of agricultural policies : a technical manual (August 1990)
Jean-Marc Burniaux, François Delorme, Ian Lienert and John P. Martin

83. Exchange Rate Policy in Advanced Commodity-Exporting Countries: The Case of Australia and New Zealand
Adrian Blundell-Wignall
Robert G. Gregory

82. Economies and the environment: a survey of issues and policy options (July 1990)
Jon Nicolaisen and Peter Hoeller

81. Financial liberalisation and consumption smoothing (forthcoming)
Adrian Blundell-Wignall, Frank Browne and Stefano Cavaglia

80. Fiscal indicators (April 1990)
Edward E. Gramlich, The University of Michigan

79. Suggestions for a new set of fiscal indicators (April 1990)
Olivier Blanchard (MIT and NBER)

78. Indicators of fiscal policy: a re-examination (April 1990)
Jean-Claude Chouraqui, Robert P. Hagemann and Nicola Sartor

77. Monetary policy in the wake of financial liberalisation (April 1990)
Adrian Blundell-Wignall, Frank Browne and Paolo Manasse

76. The dollar position of the non-U.S. private sector, portfolio effects, and the exchange rate of the dollar (February 1990)
Bixio Barenco

75. Measuring industrial subsidies: some conceptual issue (February 1990)
Neil Bruce

74. Industrial subsidies in the OECD economies (January 1990)
Robert Ford and Wim Suyker

73. The saving behaviour of Japanese households (January 1990)
Kenichi Kawasaki

72. The role of indicators in structural surveillance (January 1990)

71. Modelling business sector supply for the smaller OECD countries (October 1989)
Raymond Torres, Peter Jarrett and Wim Suyker

70. On the sequencing of structural reforms (September 1989)
Sebastian Edwards

69. The information content of the terms structure of interest rates: theory and practice (September 1989)
Frank Browne and Paolo Manasse

68. The impact of increased government saving on the economy (June 1989)
Richard Herd

67. Saving trends and behaviour in OECD countries (June 1989)
Andrew Dean, Martine Durand, John Fallon and Peter Hoeller

66. Potential output in the seven major OECD countries (May 1989)
Raymond Torres and John P. Martin

65. A post mortem on OECD short-term projections from 1982 to 1987 (February 1989)
Byron Ballis

64. Revised groupings for non-OECD countries in OECD's macroeconomic model INTERLINK (January 1989)
Paul O'Brien, Laure Meuro and Arthur Camilleri

63. Modelling housing investment for seven major OECD countries (December 1988)
Thomas Egebo and Ian Lienert

62. The economic dynamics of an ageing population: the case of four OECD economies (January 1989)
Alan J. Auerbach, Laurence J. Kotlikoff, Robert P. Hagemann and Giuseppe Nicoletti

61. Ageing populations: implications for public finances (January 1989)
Robert P. Hagemann and Giuseppe Nicoletti

60. Compatible trade and production data base: 1970-1985 (November 1988)
Claude Berthet-Bondet, Derek Blades and Annie Pin