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The "Puzzle" of Wage Moderation in the 1980s

Pierre Poret

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by

Pierre Poret
General Economics Division

November 1990



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In the 1980s, real consumption wages fell relative to labour efficiency in OECD countries, although there has been some pressure on wages recently. The persistent moderation of wages was due to labour-market slack, as well as rising non-wage labour costs and slower increases in output prices relative to consumer prices in some countries. It does not appear to reflect changes in wage behaviour. The analysis of time-series properties suggests that, typically, wages do not fully catch up to past losses. There is, furthermore, no strong evidence for instability of well-specified wage equations with the exception of France where stringent monetary policy since 1983 may have modified the formation of expectations.

* * * * *

Dans les années 80, le pouvoir d'achat des salaires a baissé par rapport à la productivité (corrigée des effets de substitution capital-travail) dans les pays de l'OCDE malgré les récentes pressions salariales. La modération persistante des salaires a été due aux déséquilibres du marché du travail, aussi bien qu'à l'augmentation des coûts indirects du travail et à des prix de production croissants moins vite que le prix de la consommation dans certains pays. Elle ne semble pas refléter un changement dans le comportement des salaires au cours du cycle. L'analyse de séries temporelles longues suggère que, typiquement, les salaires ne rattrappent pas complètement les "pertes" passées. Il n'y a pas non plus de preuves empiriques de l'instabilité d'équations de salaires bien spécifiées, sauf pour la France où le durcissement de la politique monétaire depuis 1983 aurait modifié la formation des anticipations.

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The author would like to acknowledge comments and suggestions by Andrew Dean, Robert Ford, Peter Hoeller, Jeff Shafer and participants at an internal seminar. He is also grateful to Guiseppe Nicoletti for introducing him to the econometrics of cointegration and to Valérie Barca, Anick Lotrous and Chantal Nicq for efficient research assistance.

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THE "PUZZLE" OF WAGE MODERATION IN THE 1980s

I. INTRODUCTION

Wage behaviour in the OECD area over the last decade can be characterised by greater moderation of wage demands by previous cyclical standards, despite stronger pressures in the late 1980s. However, this overall picture masks disparities across countries. Furthermore, the interpretation of wage developments is sensitive to the definition for wages, prices and productivity.

This paper investigates the extent to which wage moderation in the 1980s is a "puzzle". Whether wage developments can be reasonably well "explained" is important for forecasting, in particular for assessing whether wage moderation which persists for several years increases the risk of "catch-up" pressures emerging. The question also has implications for the effectiveness of government policies, as market-oriented reforms and tight macroeconomic policies may have led to changes in behaviour. In this paper, a change in behaviour is defined in a somewhat restrictive way as a change in the responsiveness of wages to their cyclical determinants. The model used is one in which wages are assumed to fully adjust to prices and labour efficiency in the long run and the "natural" rate of unemployment is measured by the low-frequency component of the actual rate. The extent to which structural policies may have modified the "natural" rate of unemployment is not examined; this issue was addressed in an earlier OECD study (Chan-Lee *et al.*, 1987).

Section II describes wage pressures in the 1980s, measured as the gap between real consumption wages and labour efficiency. Section III reviews the main factors influencing wage developments. It uses new measures of the natural rate of unemployment and inflation expectations which are discussed in detail in Annex A. Section IV asks whether wage moderation reflects a change in behaviour. A cointegration approach is used to assess whether there exists a stationary normal level to which wages should be expected to return in the long run. The stability of well-specified dynamic wage equations is then examined. The technical aspects of this section are presented in Annexes B and C. Some conclusions are drawn in Section V.

II. MEASURING WAGE PRESSURE

A. Defining wage pressure

Wage pressures arising from the labour market are usually measured using wage rates rather than total compensation. The wage rate is defined here as business-sector wages and salaries per employee. Wage pressures are often analysed in nominal terms. However, increases in nominal wages may only compensate for consumer price inflation, implying no income gain in real terms. They can also reflect an improvement in productivity. Equilibrium conditions require that the growth rate of wages is equal to the sum of the rates of change of prices and productivity in the long run, as the labour share would otherwise fall or increase indefinitely, other things being equal. Hence, the evolution of the ratio of real wages to labour productivity (output per employed) is used in this paper as an indicator of the intensity of wage pressures (moderation).

However, labour productivity may be determined, in turn, by wages. If adjustment lags of labour demand are not too long and the substitution elasticity between production factors is high enough, firms will reduce an excess of wages over productivity by substituting capital for labour. Labour productivity will increase and, ex post, the wage gap will be smaller than it was ex ante. Consequently, productivity has been corrected for the effect of factor substitution, so as to reflect only labour efficiency gains (1).

B. Wage moderation in the 1980s

The ratio of real wages (excluding employers' social-security contributions and using the private consumption deflator) to current labour efficiency is shown in Chart 1. This ratio has been normalised by dividing through by the historical average value over the last two or three decades (depending on data availability). Except for Switzerland and Greece, the ratio fell in all OECD countries in the first half of the 1980s, although the magnitude of the fall varied considerably between countries. In Japan and the major continental European countries, wage moderation began in the mid-1970s, largely reflecting the recovery of productivity following the OPEC I recession.

In North America, Australia and New Zealand, the fall took place only after the 1982 recession.

In the second half of the 1980s, wage developments were more contrasted across countries. Real wages have continued to decline, relative to productivity, in France, Italy, Australia, Belgium, Finland, Ireland and Spain; they have stabilised in Japan, the Netherlands, New Zealand and to some extent in the United States and Germany; but they have increased again in recent years in the United Kingdom, Canada and a number of Nordic countries.

However, as overtime and bonus payments -- which are related to cyclical movements in output per worker -- generally represent a small share of wages, real wages are unlikely to closely follow short-term fluctuations in productivity. Therefore, as productivity moves pro-cyclically, the ratio of real wages to current labour efficiency may tend, other things being equal, to increase in recession and to fall in recovery. When based on trend (i.e. cyclically-adjusted) productivity figures (2), recent wage pressures appear to be somewhat larger after removing cyclical productivity gains, with the exception of Norway. The overall picture of wage moderation in the 1980s is not significantly modified, however. In 1989, in all OECD countries except the United Kingdom, Canada, Austria, Greece, Switzerland and New Zealand, real consumption wages relative to trend productivity remained well below their historical average levels (even though a complete peak-to-peak cycle may have just been achieved).

III. STYLISTED FACTORS OF WAGE MODERATION

A. The wedge on labour

Developments in real wages are influenced by changes in output prices relative to consumer prices and by non-wage labour costs. While the real consumption wage is the relevant variable from the wage-earners' point of view, the real product cost of labour is what matters for employers. The larger the wedge between the two variables, the larger the downward pressure exerted by

employers on the wage rate. The equilibrium point finally achieved depends on the slopes of the labour supply and demand curves.

Chart 1 also displays the ratio of compensation per employee deflated by the business-sector GDP deflator ("real product compensation") to current labour efficiency. This ratio is equivalent to the usual labour-share concept except that productivity is adjusted for changes in the capital-labour ratio. In Germany, France, Belgium, Ireland, Denmark and Sweden, moderate wage increases in the 1980s could be partly related to the fact that real labour costs had moved to being above average in the first half of the decade. In the United States and Japan, wage moderation has been just sufficient to stabilise real labour costs, although at relatively high levels.

B. Inflation expectations

In setting their wage demands, wage-earners take into account expected inflation. If expectations are, for instance, below actual inflation, wage demands will be weaker than they would have been in the absence of expectation errors. The relevance of price expectations for wage behaviour, however, depends on the length of labour contracts and the frequency of cost-of-living indexation clauses.

As a first indicator, expected inflation has been proxied by the rate of change in smoothed consumer prices using the Hodrick-Prescott filter (see Part 1 of Annex A). Thus the high-frequency movements of inflation due to the oil price shocks were largely eliminated. Also, such an indicator yields expectations which are unbiased and forward-looking. With rising inflation following OPEC II assumed not to fully pass through to price expectations, expected real wages increased faster than actually observed in the early 1980s (Chart 2). In the continental European countries, perceived real wages may have been higher than actual real wages during the whole of the first half of the 1980s. On the other hand, increases in expected real wages appear to have been weaker than registered in the United States from 1982 to 1986. This result -- which is consistent with data from the Michigan consumer survey -- may partly explain why wages in the United States in the mid-1980s have not been as moderate as in the other countries. In the second half of the 1980s,

however, with inflation rates less volatile, expectational errors were small in most countries.

The picture is very different if expectations are assumed to be adaptive, that is, formed only on the basis of the observation of past inflation rates. Such a proxy implies that, in the context of the disinflation experienced from 1982 to 1988, expected inflation would have been systematically above actual inflation (Chart 2) and would not have contributed to wage moderation.

C. Labour-market conditions

Labour-market conditions also explain the divergence of real wages from labour efficiency. Data suggest, however, that for most countries there is no drift in the ratio of real wages to labour efficiency in the long run. Hence, an appropriate measure of labour-market conditions is the deviation of the current unemployment rate from some persistent or "natural" rate of unemployment, calculated in a way that ensures that the resulting unemployment gap is closed in the long run. A proxy for the natural rate has been constructed by smoothing the unemployment rate series using the Hodrick-Prescott filter (see Part 2 of Annex A). The conventional view is that the actual unemployment rate returns to the natural rate, while the alternative view -- the so-called "hysteresis" hypothesis -- assumes that the natural rate follows the actual rate. However, the filtering procedure cannot discriminate between these two assumptions and therefore this issue is not addressed here.

In most OECD countries, the 1982 recession resulted in a large gap between current and trend unemployment (Chart 3). The gap widened to a much larger extent than in the preceding trough and thus took some time to narrow despite the length of the 1980s recovery. In France and Italy, labour-market slack had still not been entirely eliminated by 1989. And the unemployment rate has only recently fallen below trend in the other major OECD countries. However, labour-market tensions have so far remained weaker than they were during the previous 1979 peak, except in Japan, Canada, Finland, Sweden and Switzerland. The ratio of job vacancies to employment provides a broadly similar picture, though indicating tighter labour-market conditions in recent

years for Japan and Austria than suggested by the estimated unemployment gaps (Table 1).

A quantitative analysis of factor contributions to wage developments, based on estimated wage equations using unbiased inflation expectations, is set out in Annex C (Part 2). The results for the major seven countries confirm that none of the factors just reviewed -- the wedge on labour, inflation expectations, labour-market conditions -- has exerted a positive influence on the gap between real consumption wages and labour efficiency in the 1980s on average, except for relative output prices for Germany and the employers contributions for the United Kingdom (see Table C2 in Annex C). For 1980 to 1989, the equations predict a real wage growth which is lower than productivity growth by 0.2 percentage points per annum in North America and the United Kingdom, 0.4 in Germany, 0.6 in Sweden, 0.7 in Japan, 0.8 in France, and more than one point in Italy, Australia and Spain.

IV. HAS AGGREGATE WAGE BEHAVIOUR CHANGED?

In this section, the issue of whether wage behaviour has been atypical over the current cycle is addressed in two ways. First, the "puzzle" of wages being persistently below average is examined by challenging the existence of a stable long-run norm of real wages relative to productivity levels. Second, the stability of well-specified dynamic equations is tested to assess whether wage behaviour has changed.

A. Are wages determined in relation to a stable long-run level?

The hypothesis that wages tend to revert to some normal level in the long run has been rejected using a cointegration approach (see Annex B). Only for the United Kingdom was a stationary combination between the level of wages and its long-run determinants found at usual confidence levels. For the other countries, therefore, wage "losses" or "gains" are never entirely recovered (or, as the tests may have low power against long-memory stationary processes, the time needed to recover them may be very long). These results hold even when the 1980s are removed from the estimation sample period.

The implications are two-fold. First, the observation that real wages relative to productivity have remained below the historical average in most OECD countries over the 1980s cycle does not indicate per se a break in wage behaviour, as typically developments in real wage levels are characterised by persistence. Second, while the gap between real wages and productivity is found to be non-stationary by historical standards, its first difference will be stationary and therefore the stability of wage behaviour may be tested using equations specified in growth-rate terms.

B. Stability of wage equations

Section III has reviewed the main identifiable factors influencing wage developments in the 1980s. To assess whether large residuals are left unexplained, however, a stability analysis of wage equations is needed. The equation specification (presented in more detail in Part 1 of Annex C) is characterised by three important features: i) the elasticities of wages with respect to prices and productivity are unity, thereby ensuring well-behaved long-run properties of the model; ii) inflation expectations are assumed to be unbiased and forward-looking, as proxied using the Hodrick-Prescott filter; iii) the "natural" rate is measured by the smoothed rate of unemployment and explicitly enters the model instead of being derived from it, as is often done.

Stability diagnostics are conditional on model specification. The model used in this paper is not suitable for isolating the impact on wage demands of structural reforms aimed at enhancing trend labour productivity and reducing the natural rate of unemployment. These effects are, in principle, already captured in the right-hand side variables of the model. Therefore, the model is likely to be inherently more stable than a specification which omits such important factors. On the other hand, it is appropriate to check the stability of the sensitivity of wages to their determinants.

Recent studies for the United States (see Adams and Coe (1990) and references therein) found a significant tendency of equations to overpredict wage increases in the 1980s. However, no such a tendency can be detected in the U.S. equation reported here: residuals are relatively small and evenly balanced (Table C1 and Chart B in Annex C). Residuals are larger for the other

countries but Chow stability tests do not indicate that the tracking performance of the equations is significantly different in the 1980s compared with earlier periods.

Nevertheless, closer scrutiny of residuals indicates a tendency towards underprediction in Canada and Italy and overprediction in Germany, France and Australia. To better assess the significance of these tendencies, dummy variables set at unity from 1980 and 1985 have been successively tested. Their coefficients were found to be significant for France and, to a lesser extent, Italy -- with negative and positive signs, respectively (Table C1 in Annex C). Some studies -- but not all -- have found that equations for Australia tend to overpredict wage increases since the 1983 Prices and Income Accord, with dummy variables set at unity from 1983 being significant (see Chapman and Gruen (1989) for a review of Australian studies). However, such dummies are not significant in the equation presented in this paper.

C. Aggregation problems and changes in expectations formation

For Italy, the under-estimation of wage increases may pertain to an aggregation problem (3). As the North and Centre regions represent the largest share of the national wage bill and employment, the unemployment gap prevailing in these areas is likely to be the relevant indicator for aggregate wage behaviour (see the April 1990 OECD Economic Survey of Italy). Until 1983, the aggregate unemployment rate moved closely in line with the unemployment rate of the North and Centre regions; but afterwards, the former continued to steadily increase while at the same time the latter stabilised and then fell, suggesting that the labour market was tighter in the second half of the 1980s than indicated by the aggregate unemployment gap. Due to data limitations, however, this aggregation problem hypothesis has not been formally tested.

For France, a recent study (Ralle and Toujas-Bernate, 1990) attributes the overprediction to a price elasticity of wages below unity after 1982. Long-lasting under-indexation is however implausible. Rather, systematic negative residuals disappear when allowance is made for weaker inflation expectations. Price expectations, indeed, are found to mainly rely on past inflation rates before 1983 and to be more forward-looking afterwards (see

Part 3 of Annex C). In the French context of disinflation, the shift from adaptive to unbiased expectations implies (by construction) a reduction in expected inflation and is estimated to have accounted for one percentage point per annum in the deceleration of wage increases from 1983 to 1987.

The finding of significant shifts in price parameters for the French wage equation gives some support to the well-known Lucas (1976) critique according to which perceived changes in macroeconomic policy regimes make econometric reduced-form relationships unstable as expectations change with policies. Such results for France can therefore be viewed as a positive test that disinflationary policy has gained credibility since 1983. Other studies using different testing approaches for credibility yielded the same conclusion (Giavazzi and Giovanni, 1989; Artis and Nachane, 1990). Lower inflation expectations can be related to the strategy of pegging the franc to the deutsche mark, thereby "importing the credibility" of the German central bank (4).

V. CONCLUSIONS

Real consumption wages have fallen relative to labour efficiency in the first half of the 1980s and have remained below historical averages in the second half of the decade in most OECD countries. Wage pressures have significantly increased only recently and in a limited number of countries (the United Kingdom and Canada among the seven major countries). This picture is consistent with the labour-market conditions prevailing in the 1980s, which were marked by a deep recession in the early 1980s and a peak in the late 1980s which was not as high as in previous cycles. In addition, wage moderation partly reflects an increase in the wedge on labour, with real product labour costs growing faster than real consumption wages.

The evidence presented here indicates that aggregate wage behaviour has in general not changed in the 1980s. First, the analysis of time-series properties over the last three decades suggests that wages have not been determined in relation to a stable long-run equilibrium level, so that real wages persistently below (above) average are not per se the sign of a break in

behaviour. This finding also implies that the observation that wage levels have departed from some "warranted" level need not, in itself, lead to catch-up. Second, there is no strong evidence for instability in the wage equations used in this paper. These equations include, in particular, proxies for changing trend productivity growth and "natural" rates among the explanatory variables. Structural policies have probably influenced such variables. However, it does not appear that they have led to increased responsiveness of wages to cyclical labour-market slack (as indicated by gaps between the actual rate and "natural" rate of unemployment), nor has tighter macroeconomic policy caused a shift in the price parameters in wage equations. The only exception is for France for which a significant tendency towards overprediction was detected from 1983, probably reflecting greater credibility of disinflationary monetary policy.

NOTES

1. Labour efficiency is calculated assuming that technical progress is only labour-augmenting in a three-factor Cobb-Douglas production function with constant returns to scale: $\ln Q = (1-a-b) \ln K + a \ln [L \cdot E] + b \ln EN$, where Q is output, K capital, L labour, E labour efficiency, EN energy, a and b are the (sample period average) labour share and energy share in business-sector GDP, respectively. ln denotes the logarithm operator. For the smaller countries, energy is excluded.
2. Trend labour efficiency has been calculated by applying the Hodrick-Prescott filter (see Annex A).
3. Measurement errors in the variables are also part of the explanation for equation instability, as non-wage labour costs are OECD Secretariat estimates based on old National Accounts figures (i.e. before rebasing).
4. Some evidence for a specific EMS effect for other European countries is provided in Giavazzi and Giovanni (1989), Artis and Nachane (1990), and Kremers (1990). An interesting parallel can be made with other large monetary regime shifts such as the abandonment of the gold standard in 1914 which Alogoskoufis and Smith (1989) found to cause significant shifts in the price parameters of wage equations for the United States and the United Kingdom.

Table 1

Vacancy rates and unemployment gaps for selected years

	Trough				Peak			
	1975*		1983*		1979*		1989*	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
United States	1.9	76.2	2.1	75.7	-1.4	127.0	-1.0	100.0
Japan	0.2	86.4	0.2	84.5	-0.2	93.9	-0.3	133.8
Germany	0.9	107.5	1.5	40.9	-0.8	138.2	-0.6	111.7
France	0.0	113.8	1.0	50.0	-0.5	92.1	0.1	81.2
Italy	-0.1	..	0.4	..	-0.6	..	0.3	..
United Kingdom	0.1	51.2	2.2	83.8	-2.2	135.2	-1.5	126.1
Canada	0.2	120.0	2.2	54.1	-0.9	107.6	-1.2	122.7
Australia	0.7	..	2.4	46.9	-0.4	86.0	-1.4	145.3
Austria	0.0	99.8	0.8	48.7	-0.7	113.6	-0.1	143.9
Denmark	0.9	60.2	1.8	22.4	-0.7	116.0	-0.5	66.6
Finland	0.2	77.1	0.3	77.5	-0.2	84.2	-1.3	213.8
Netherlands	0.4	116.4	3.2	24.9	-1.9	166.4	-0.6	..
Norway	0.6	86.7	0.9	45.3	-0.5	110.0	-0.3	114.1
Spain	-0.6	..	2.0	..	-1.8	..	-0.5	..
Sweden	0.3	87.7	0.8	53.6	-0.3	136.7	-0.4	127.2
Switzerland	0.1	33.8	0.0	101.5	-0.2	146.2	-0.7	198.6

* Or closest year.

(1) Difference between current and trend unemployment rate (as percentage rate point).

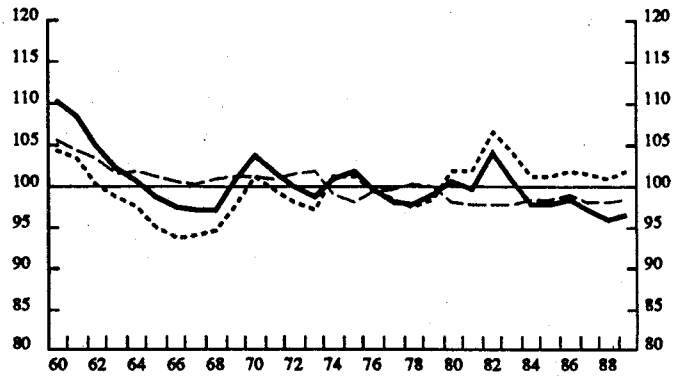
(2) Ratio of job vacancies to business-sector employment. 100 = 1973-1989.

Source: OECD, Main Economic Indicators.

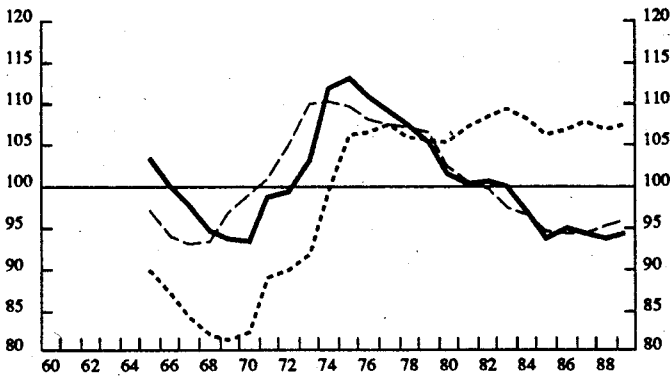
Chart 1. Indicators of wage pressure
100 = sample-period average

- Real consumption wages/labour efficiency
- - - Real product compensation/labour efficiency
- · - Real consumption wages/trend labour efficiency

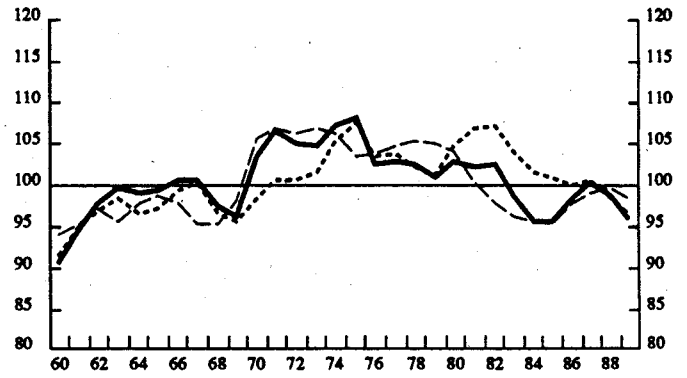
United States



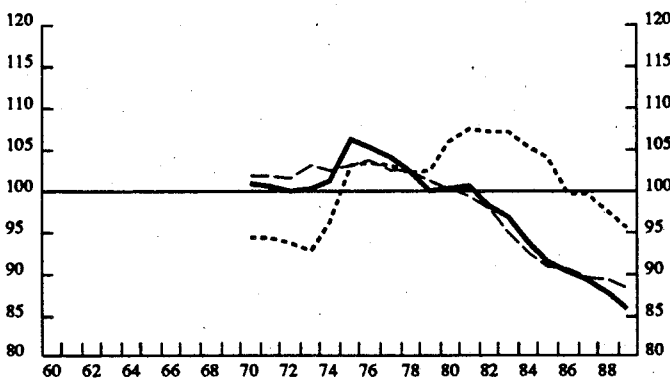
Japan



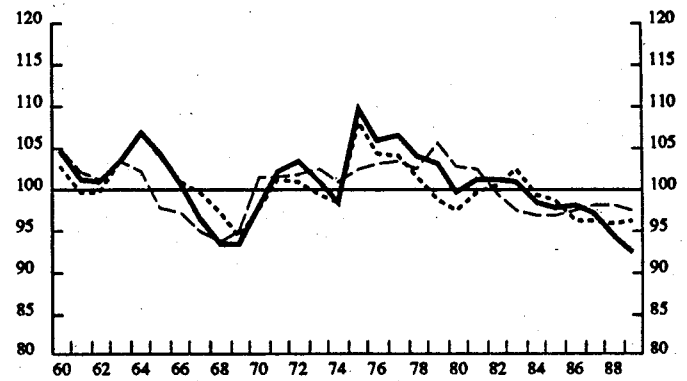
Germany



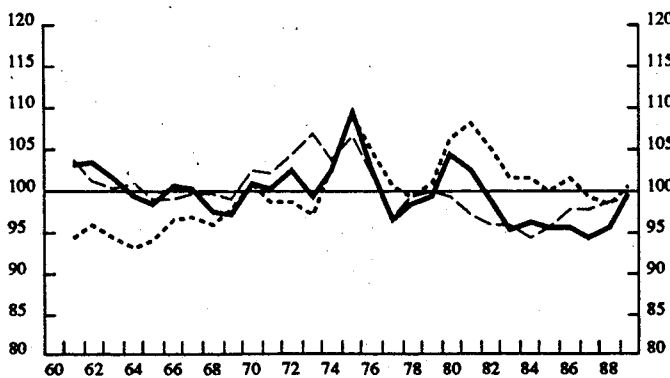
France



Italy



United Kingdom



Canada

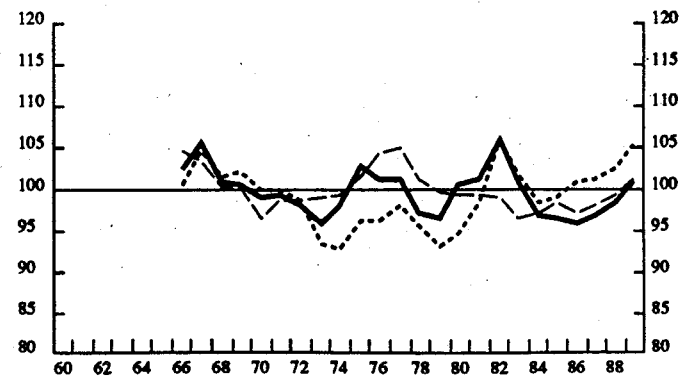
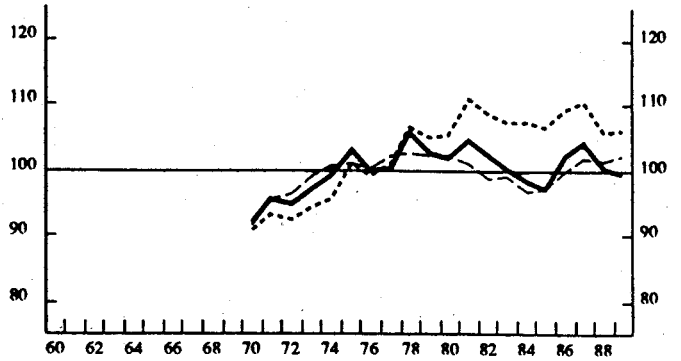


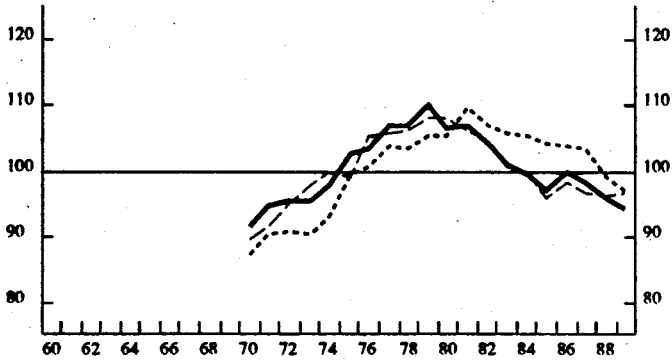
Chart 1 (cont.). Indicators of wage pressure
100 = sample-period average

- Real consumption wages/labour efficiency
- - - Real product compensation/labour efficiency
- · - Real consumption wages/trend labour efficiency

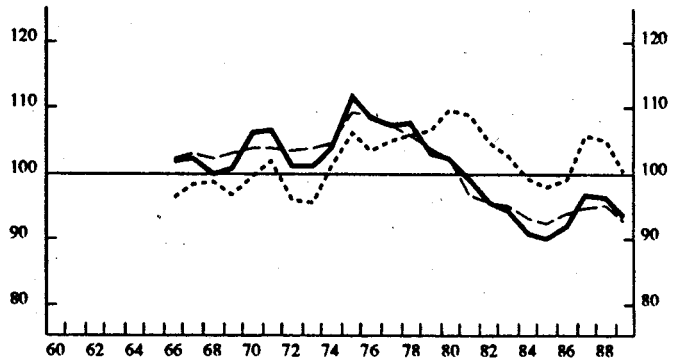
Austria



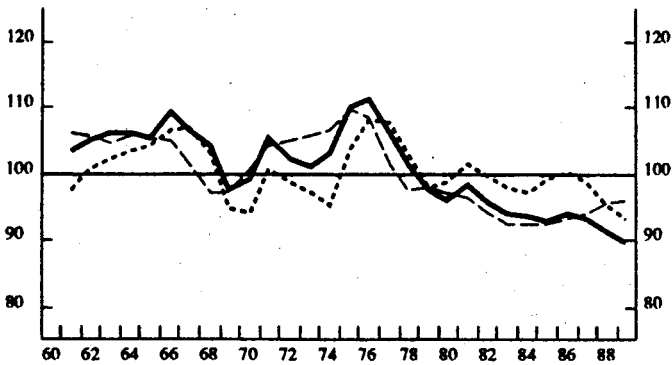
Belgium



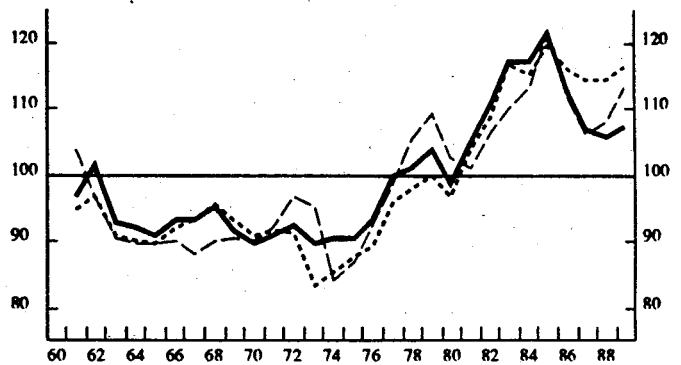
Denmark



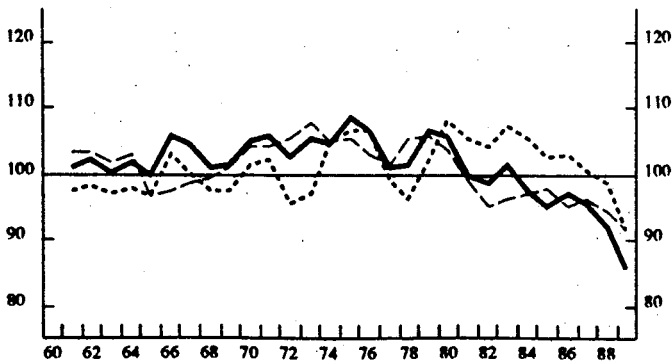
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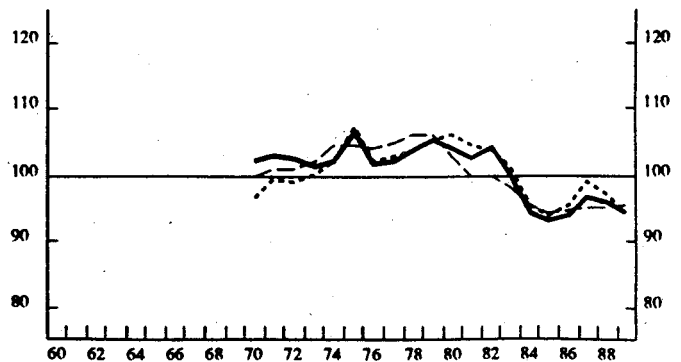
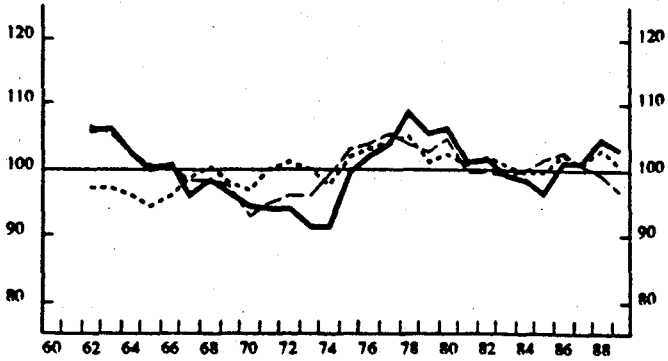


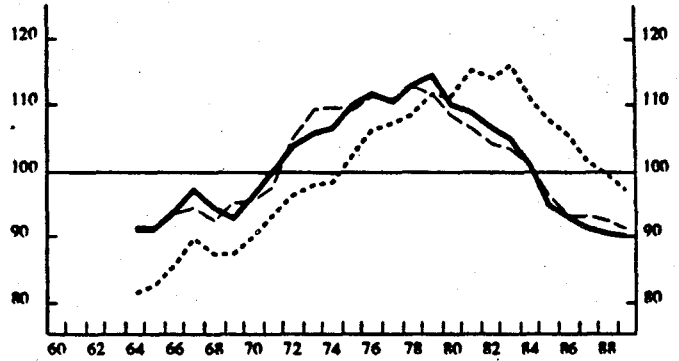
Chart 1 (cont.). Indicators of wage pressure
 100 = sample-period average

- Real consumption wages/labour efficiency
- Real product compensation/labour efficiency
- - - Real consumption wages/trend labour efficiency

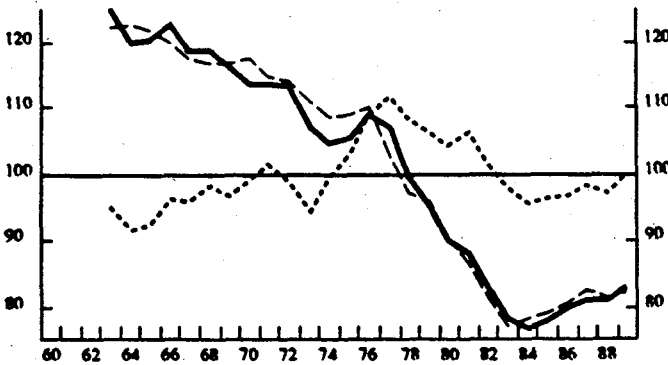
Norway



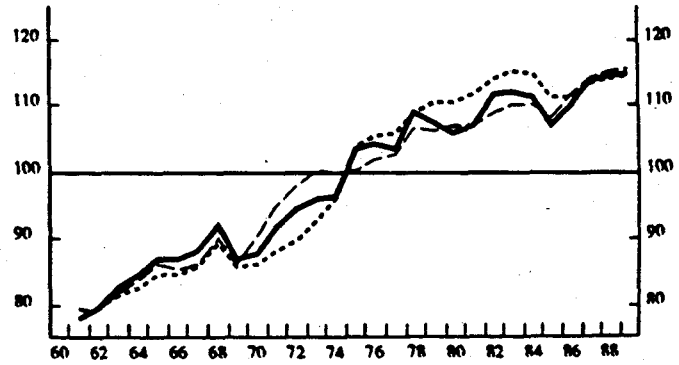
Spain



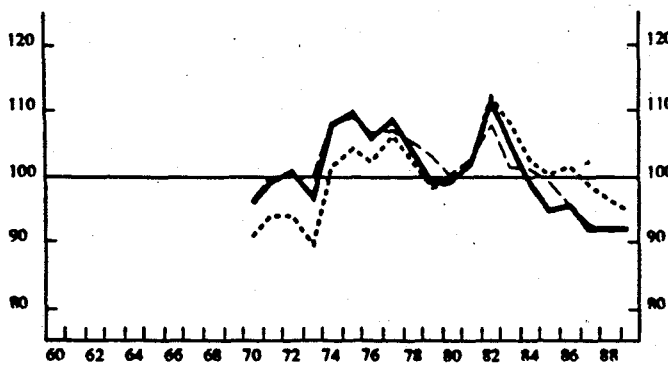
Sweden



Switzerland



Australia



New Zealand

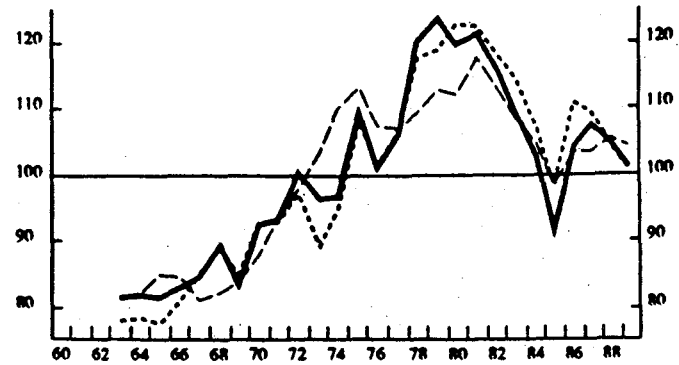
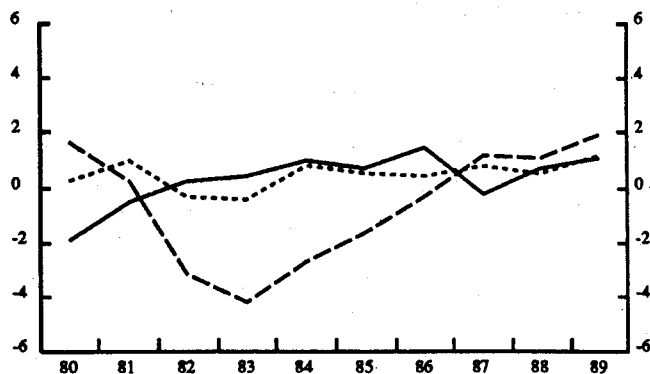


Chart 2
Actual and expected real consumption wages

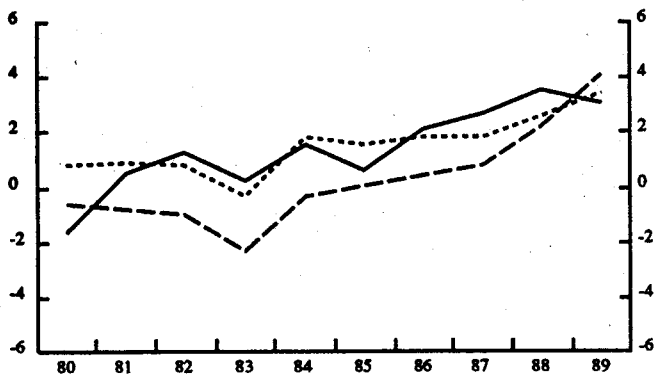
Percentage changes in wages less:

- Current inflation
- Trend inflation
- - - Four-semester moving average past inflation

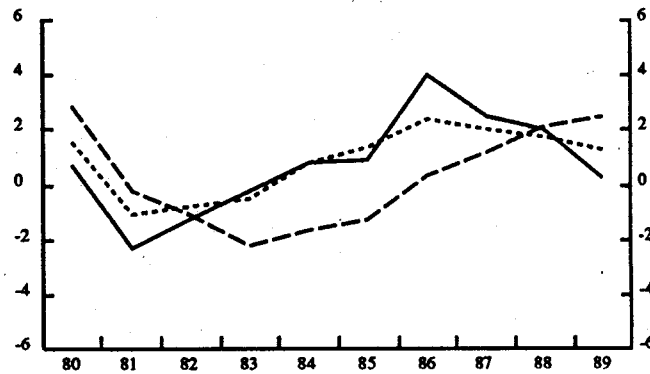
United States



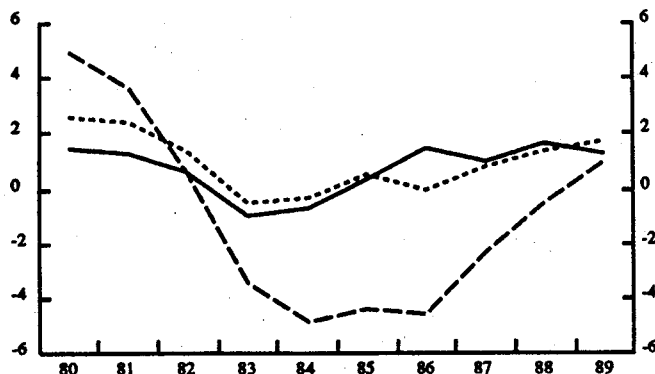
Japan



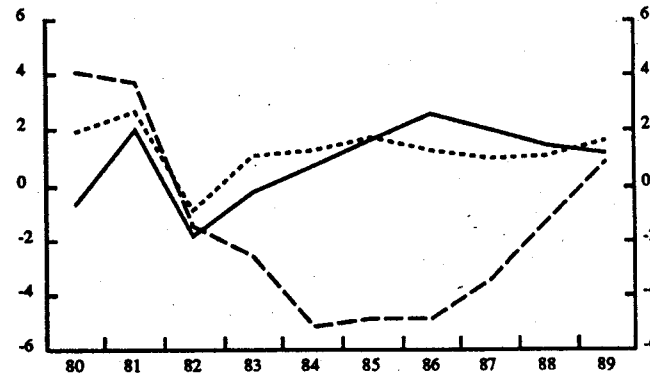
Germany



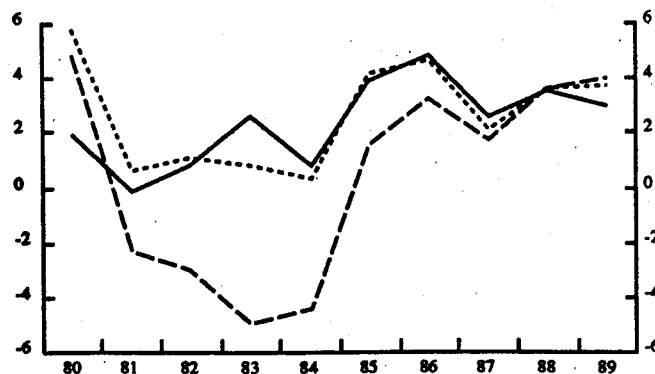
France



Italy



United Kingdom



Canada

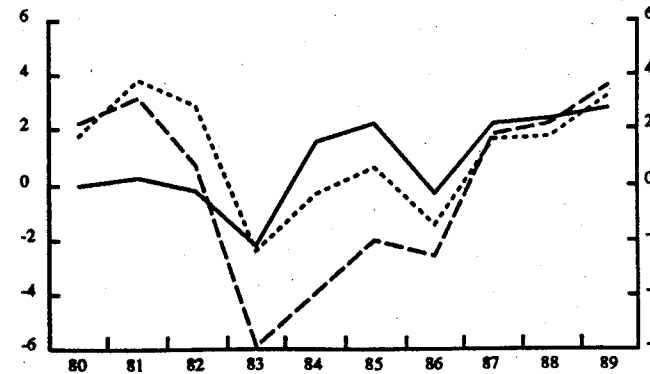
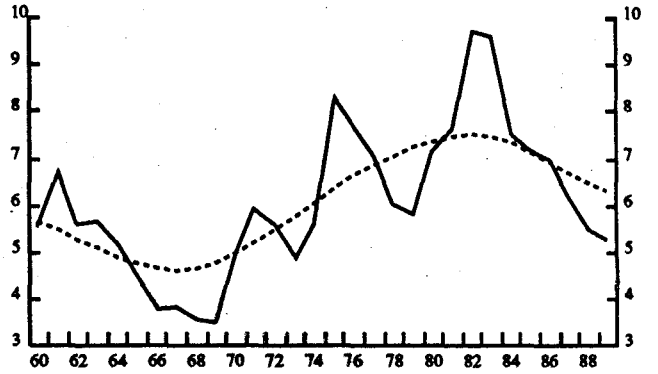


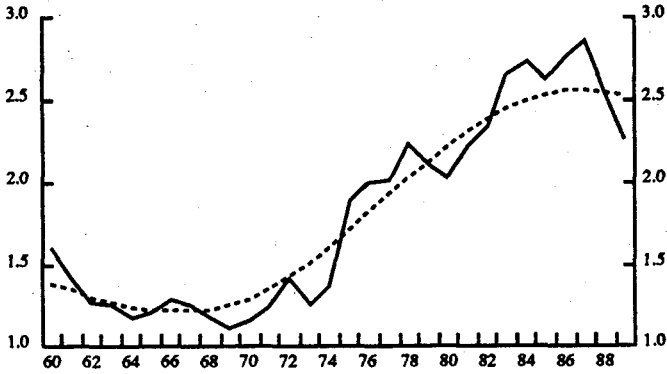
Chart 3
Actual rate of unemployment and estimated natural rate
 (per cent of labour force)

— Current unemployment
 Trend unemployment

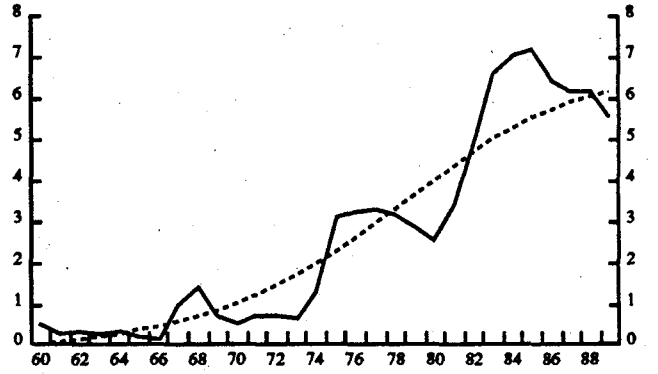
United States



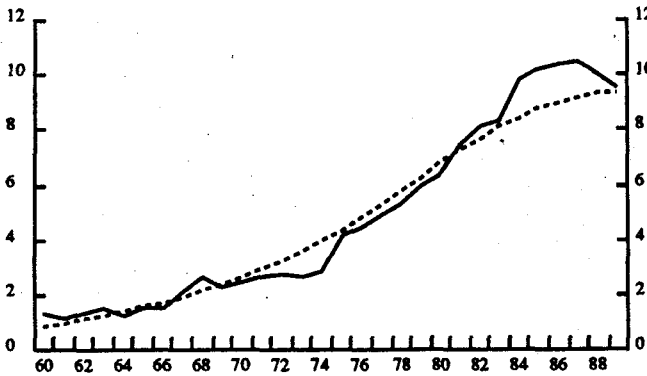
Japan



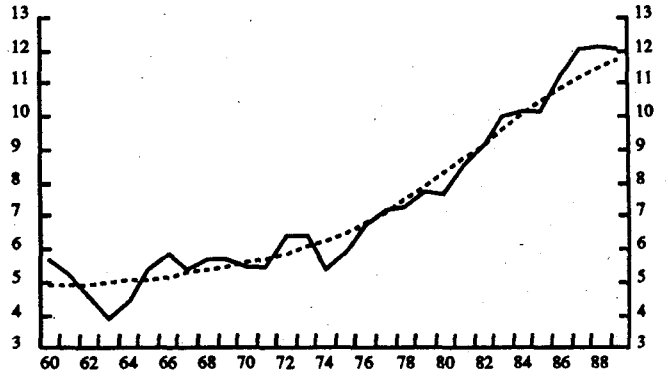
Germany



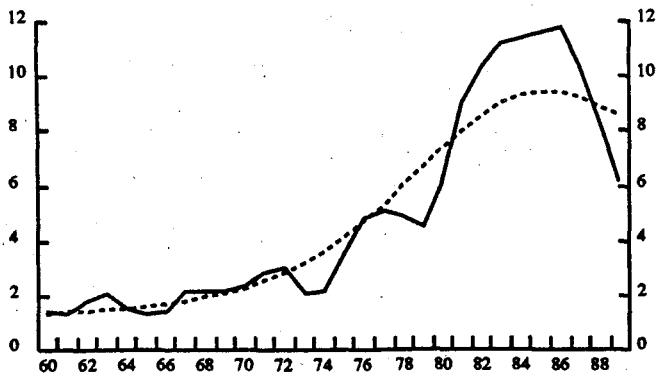
France



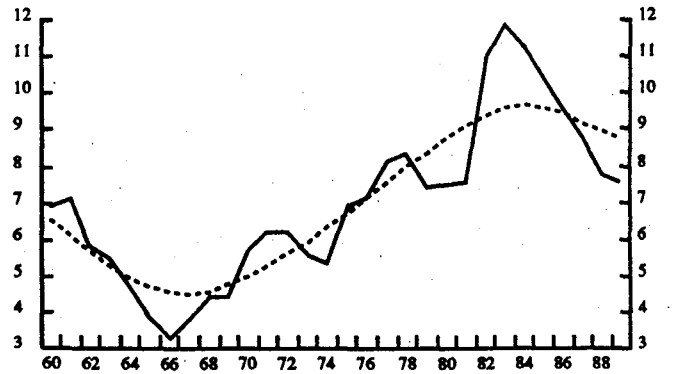
Italy



United Kingdom



Canada



ANNEX AINDICATORS OF INFLATION EXPECTATIONS AND THE NATURAL RATE OF UNEMPLOYMENT

This annex examines the rationale behind the use of the Hodrick-Prescott filter to measure inflation expectations and the natural rate of unemployment. This filter, which is increasingly used in economic research, in particular in Real Business Cycle theory, 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 from the series under the constraint that the sum of the squared second differences does not exceed a certain factor, which is chosen by the user. It can be interpreted as a time-varying trend. In extracting the low-frequency component from the series, this filter uses backward and forward observations. In this paper, in order to get reliable figures for the late 1980s, the filter used series extended by the June 1990 OECD Economic Outlook and internal medium-term projections.

The critical aspect of the procedure is that the split of the series into "transitory" and "permanent" components is a matter a judgement. A very small smoothing factor, for instance, implies that most of the shocks to the series are changes in the trend, while a very large factor leads to an almost constant trend with shocks being assumed to be mainly cyclical. In this study, the smoothing factors were first selected uniformly across countries by visually checking the "plausibility" of the estimated trends. In view of regression results using alternative smoothing factors, they were then somewhat increased for the unemployment-rate variable for Germany and France, and the price variable for Canada.

1. Inflation expectations

Price expectations are commonly assumed to be adaptive and are modelled as a distribution of past and current inflation rates. However, for most time-series processes commonly observed, this assumption implies that expectations are sub-optimal (they are biased and do not minimise expectational errors given available information).

A growing body of literature assumes more forward-looking expectations. Proponents of the Rational Expectation Hypothesis use one-period ahead inflation forecasts from reduced-form price equations or from complete macroeconomic models. Although this approach has proven successful in explaining wage behaviour (McCallum, 1976; Atesoglu, 1988; Moghadam and Wren-Lewis, 1989; Blanchard and Sevestre, 1989), it has not been demonstrated to have significantly greater predictive power than adaptive-expectation models (Ormerod, 1982; Coe, 1985; Moghadam and Wren-Lewis, op. cit.): Another measure of forward-looking expectations is the inflation rate implicit in the term structure of interest rates. This indicator has been used in Paunio and Suvanto (1981) in wage equations using Finnish data. However, the extent to which inflation expectations from financial markets are relevant for the labour market is debatable. A recent study by Englander and Stone (1989) shows that forecasts from financial market surveys contain little information for wage determination. A final possibility is to use survey data as a direct measure of expectations. Englander and Stone (op. cit.) found that the U.S. Michigan household survey is a good predictor of future inflation and outperforms adaptive expectations in wage equations.

This paper uses yet another method to extract expected inflation from the data. Consumer prices were smoothed using the Hodrick-Prescott filter, on the a priori assumption that only the low-frequency component of inflation is anticipated. The use of this filter implies, for example, that the contribution of an oil shock to trend inflation, if any, is assumed to have been anticipated while its transitory effect on inflation has been filtered out. In practice, however, the selected smoothing factors were such that oil shocks affected the trend inflation rates only marginally.

By construction, this approach yields unbiased and forward-looking expectations, as the filter uses backward and forward observations. However, it does not guarantee that expectations are "rational", since the expectational errors are likely to be correlated with publicly available information. In that sense, the less auto-correlated the errors are, the "more" rational are the expectations. The degree of serial auto-correlation depends on the time-series properties of inflation rates and the smoothing factor. With the smoothing factor chosen, expectational errors were by no means persistent, with the auto-correlation coefficients falling rapidly (Table A1). But the first-order auto-correlation coefficient is not zero, averaging 0.4 across countries. Results of wage equation estimates (see Annex C) were not improved when error persistence was further reduced (by lowering the degree of smoothing). This does not imply that a certain dose of "irrationality" in expectations is needed to explain wage behaviour, as rational expectations cannot be generated with this filter other than in the extreme case of perfect expectations (obtained by using very small smoothing factor).

As a rough test, wage equations were estimated using the smoothed inflation rate and, as an alternative, a four-semester moving average of past inflation. For almost all countries, the standard errors of estimates using smoothed inflation were somewhat smaller. For the United States, there was virtually no difference. For France, standard errors were smaller only if the moving-average expectations were used before 1983 and the smoothed rate afterwards (see Part 3 of Annex C). On the whole, the conclusion is that the forward-looking smoothing hypothesis is at least as good as the adaptive approach.

2. The natural rate of unemployment

The "natural" rate of unemployment is the rate which prevails when the labour market is at equilibrium. It can be defined as the level at which the actual unemployment rate ceases to exert any downward or upward pressures on real wages relative to labour efficiency. The aggregate natural rate is likely to vary with a whole range of factors including the generosity and duration of replacement income, the costs of adjustment between supply and demand for

labour, and changes in the composition of the labour force (since the natural rate can differ according to sex, age, profession and region).

A direct way of measuring the natural rate is to estimate an equation for the current rate of unemployment, using such structural factors as well as "cyclical" variables as the right-hand side variables. An alternative -- but theoretically equivalent -- strategy is to estimate a wage equation where these factors are entered in addition to the current rate of unemployment and other standard variables. In both cases, the natural rate is then calculated as the sum of the estimated contributions of the structural factors.

In this paper, the "natural" component of unemployment has been estimated using the Hodrick-Prescott filter, on the grounds that factors affecting that natural rate are infrequent, slow moving and persistent. The computational ease of this filter is also an advantage of this approach. Unit-root tests (not reported here) suggest that the unemployment rate is non-stationary (i.e. it does not revert to some average value or deterministic time trend). King and Rebelo (1989) show that the Hodrick-Prescott filter is appropriate for extracting the permanent component from a non-stationary series. There exists a critical degree of smoothing below which the resulting trend can preserve the long-run properties of the series, while the deviations from the trend are made stationary (weakly auto-correlated). This stationarity condition is necessary for any proxy for labour market disequilibrium to be correlated with deviations of real wage growth from productivity growth, because the latter are also stationary. The use of a mean value or a deterministic time-trend to model the natural rate, for instance, yields non-stationary unemployment gaps and spurious correlations in regression analysis.

The results from the filtering approach are corroborated by the following pieces of evidence:

- In terms of pattern, the smoothed unemployment rates are very close to the natural rates estimated by the IMF (Adams and Coe, 1990) and the Bank of Canada (Ford and Rose, 1989) using structural models. For the United States, however, the average natural rate estimated by

the IMF is below the average unemployment rate by 0.5 percentage points over the 1965-88 estimation period, which is inconsistent with the observation that the ratio of real wages to labour efficiency does not exhibit any long-run drift over the sample period (Table A2). On the other hand, by construction, such a bias cannot occur using the smoothed rate.

- In number of OECD countries, the smoothed rates appear to reflect so-called "frictional" unemployment rates, as measured by shifts in the Beveridge curve. The correlation of the job vacancy rate is much higher with the gap between the actual unemployment rate and the smoothed rate than it is with the unemployment rate alone, for the United States, France, the United Kingdom, Canada, Belgium, Finland, Norway and Switzerland (Table A3). Only in Germany, Austria, Denmark and the Netherlands, is the correlation weaker. A formal estimate of the natural rate from Beveridge curve shifts is reported in Table A2 for the United States, a country that probably has the most reliable vacancy data.
- As reported in Section IV, these gaps prove to have a significant impact in the wage equations for European countries, for which econometricians have typically had difficulty finding a robust influence of labour-market conditions.

However, the smoothed rates differ from other recently published OECD estimates, the so-called "non-accelerating wages rates of unemployment" (NAWRU) (Torres and Martin, 1990). Such estimates rely on estimating a Phillips curve in which the perceived natural rate and expected trend productivity growth are assumed not to adjust over time and are therefore modelled by a constant term. The Phillips curve is then solved to get the unemployment rate required to close the gap between real wage growth and some "true" growth rate of trend productivity. At best, however, this can be viewed as a sort of "short-term" natural rate, since misperceptions in trend productivity cannot plausibly be assumed to last forever.

Table A1

Consumer price inflation: autocorrelation function.

	United States	Japan	Germany	France	Italy	United Kingdom	Canada	Australia	Spain	Sweden
<u>Actual inflation</u>										
Lag: 1	0.88	0.85	0.72	0.84	0.90	0.84	0.89	0.62	0.36	0.52
2	0.74	0.63	0.70	0.69	0.78	0.71	0.76	0.59	0.52	0.57
5	0.40	0.37	0.26	0.33	0.61	0.42	0.23	-0.04	0.48	0.35
10	0.20	0.04	-0.18	-0.10	0.10	0.26	-0.13	-0.07	0.05	0.03
<u>Trend inflation</u>										
Lag: 1	0.99	1.00	0.99	0.97	1.00	0.99	0.96	0.93	0.99	0.99
2	0.95	0.97	0.95	0.91	0.98	0.96	0.90	0.82	0.96	0.96
5	0.76	0.76	0.61	0.60	0.80	0.78	0.65	0.32	0.73	0.76
10	0.27	0.21	-0.06	-0.07	0.26	0.24	-0.00	-0.20	0.02	0.16
<u>"Expectation errors"</u>										
Lag: 1	0.51	0.50	0.08	0.37	0.38	0.37	0.76	0.17	-0.42	-0.17
2	0.07	-0.13	0.17	-0.06	-0.23	-0.00	-0.50	0.30	-0.01	0.02
5	-0.44	-0.23	0.15	-0.22	-0.12	-0.34	-0.36	-0.33	0.19	-0.17
10	0.09	-0.18	-0.06	-0.01	-0.36	0.35	-0.17	0.09	0.06	-0.03

Table A2
Various indicators of natural rates for the United States
 (Per cent of labour force)

	1965-88	1965-70	1970-75	1975-80	1980-85	1986	1987	1988
Hodrick-Prescott filter	6.2	4.8	5.6	6.9	7.4	6.9	6.7	6.5
Beveridge curve shift (a)	6.3	4.4	5.7	7.4	7.3	7.3	6.9	6.1
IMF estimate (b)	5.7	4.0	5.2	6.7	6.8	6.3	6.1	5.9
INTERLINK model	6.6	6.3	6.6	6.9	6.8	6.5	6.5	6.5
Memorandum item: actual unemployment rate	6.2	4.0	5.9	7.0	8.1	7.0	6.2	5.5

a) Calculated as $\exp[\ln \text{UNR} + \ln \text{VAC}]$, where UNR is the actual unemployment rate and VAC is the ratio of help-wanted advertising (Conference Board) to total business employment. VAC has been normalised (divided by its average value) so that $\ln \text{VAC}$ is equal to zero on average over the sample period. It is assumed that the unemployment rate varies negatively with the job vacancy rate around the natural rate (UNR*) according to the following model: $\ln \text{UNR} = \ln \text{UNR}^* - \ln \text{VAC}$. Hence, the exponential of the sum of $\ln \text{UNR}$ and $\ln \text{VAC}$ provides an estimate of the natural rate. This estimate is also equal to the exponential of the sum of the constant term (a) and the residuals (ϵ) from a Beveridge curve equation as follows: $\ln \text{UNR} = a + \ln \text{VAC} + \epsilon$.

b) See Adams and Coe (1990).

Table A3

Correlation between unemployment gaps and vacancy rates (a)

	United States 1960-1989	Japan 1962-1989	Germany 1960-1989	France 1963-1989
(1)	-0.80	-0.71	-0.33	-0.70
(2)	-0.42	-0.51	-0.87	0.08
	United Kingdom 1962-1989	Canada 1966-1989	Austria 1960-1989	Belgium 1960-1989
(1)	-0.51	-0.64	-0.67	-0.34
(2)	-0.27	-0.28	-0.78	-0.14
	Denmark 1970-1988	Finland 1961-1989	Netherlands 1970-1987	Norway 1962-1989
(1)	-0.78	-0.57	-0.87	-0.67
(2)	-0.84	-0.16	-0.89	-0.40
		Sweden 1962-1989	Switzerland 1975-1989	
(1)		-0.79	-0.82	
(2)		-0.71	-0.17	

a) Vacancy rates are ratios of job vacancies to business-sector employment.

Source: OECD, Main Economic Indicators.

- (1) Using ratios of the unemployment rate to the smoothed rate.
 (2) Using ratios of the unemployment rate to its average value.

ANNEX BDO REAL WAGES RELATIVE TO PRODUCTIVITY REVERT TO TREND?

The hypothesis that wages tend to catch up to some stable "normal" level in the long run was examined for the seven major OECD countries within the cointegration framework. This consists of estimating a static relationship between the wage rate and its long-run determinants and then testing whether the effects of shocks to this relationship tend to die out over time. In practice, this amounts to testing for the stationarity of the residuals from the static regression (see the formal definition of stationarity in Table B1).

Most studies emphasise the role of insider influences and market imperfections to rationalise a specification in terms of levels (the "target wage" model). However, a specification in levels is also consistent with a market-clearing framework, the equilibrium real wage rate being determined at the intersection of the labour demand and supply curves. At this point, the real wage rate is equal to marginal productivity.

As a first step, the long-run stationarity of the ratio of real consumption wages to current labour efficiency (in logarithm) has been tested by means of unit-root tests. This is conceptually equivalent to applying a cointegration test to a regression of wages on prices and productivity in which unitary elasticities are imposed. The tests do not reject the null hypothesis that the ratio of real wages to labour efficiency is non-stationary in all the seven major countries at the 5 per cent level (Table B1). The ratio was found to be integrated of order one; that is, its first difference is stationary.

However, the failure to find a stationary ratio of real wages to labour efficiency may be due to the omission of factors that permanently shift the labour demand or supply curve. Unfortunately, data limitations greatly restrict the number of potentially cointegrating variables that can be

considered. The following were included in static regressions insofar as they were integrated of order one and their coefficients had the right sign:

- the business-sector GDP deflator relative to the consumption deflator;
- the rate of employers' contributions to social security;
- the share of women in the labour force, assumed to capture the negative effect of sex discrimination on wages;
- the ratio of working-age to total population; this demographic factor causes shifts in the aggregate labour-supply curve; and
- the average effective tax rate on personal income, which affects the equilibrium level of the before-tax wage rate.

Even enriched in this way, however, the model passes the cointegration test at the 5 per cent level only for the United Kingdom (Table B2). For North America and Germany, the equations are close to passing the test at the 10 per cent level. These overall results hold even when the estimation period does not include the 1980s.

It is worth noting, however, that consistent with the cointegration framework, the tests use actual data rather than expectation and trend variables, the latter being equal to current values at equilibrium. For the same reason, the unemployment rate variable is not included, since the labour-market gap is assumed to be the appropriate variable, and this is closed in the long run. Other studies have found cointegrating wage equations using the unemployment rate among the right-hand-side variables. However, when the actual unemployment rate was added to the equations as specified in Table B2, results (not reported here) were improved for Germany only, although the equation still did not pass the Dickey-Fuller test at the 10 per cent level. To control for simultaneity bias, the one-period lagged value of the unemployment rate was also used, but this did not change the overall results.

Table B1
Unit root tests

Variable: logarithm of the ratio of real consumption wages
to current labour efficiency
(business sector)

	Alternative 3		Alternative 2	Alternative 1
	Φ_3	Φ_2	Φ_1	$t_{\alpha'}$
United States (59)	6.02*	5.47	--	--
Japan (49)	2.13	1.91	0.87*	0.98
Germany (55)	2.75	2.20	0.99*	0.89
France (53)	1.45	2.38	2.30*	1.89
Italy (59)	2.82	1.89	3.24*	0.07
United Kingdom (57)	4.97	3.81	2.16*	0.93
Canada (47)	3.35	2.36	1.13*	0.45

In parentheses: number (T) of observations (semi-annual data). An asterisk denotes that the non-stationarity hypothesis cannot be rejected.

Description of the test

A variable, χ , will be stationary if the coefficient (ρ) of the one-period lagged χ is less than one in an equation such as: $\chi_t = \mu + \rho\chi_{t-1}$, or equivalently if $\alpha = \rho - 1$ is less than zero in $\Delta\chi_t = \mu + \alpha\chi_{t-1}$.

More specifically, the test is the Augmented Dickey-Fuller test (using fourth-order correction). The testing strategy (Perron, *Journal of Economic Dynamics and Control* (12), 1988) goes from "general" to "specific". One starts by testing the null hypothesis of a unit root with a drift and no time trend ($\Delta\chi_t = \mu + \alpha\chi_{t-1} + \epsilon_t$) against Alternative 3, which is:

$$\Delta\chi_t = \mu + \beta(t-T/2) + \alpha\chi_{t-1} + \sum_{i=1}^4 \gamma_i \Delta\chi_{t-i} + u_t$$

The test statistic (Φ_3) is therefore for the constraint that $\alpha = \beta = 0$. If Φ_3 exceeds the critical value, one rejects the null (i.e. "accepts" that χ is stationary, perhaps with a time trend; the hypothesis that $\alpha = 0$ and β is non-zero is not tested) and the procedure ends. If not, one would "accept" that χ is a random walk, but this result could be due to the fact that the assumption of a non-zero drift in the null was erroneous. To check that, one repeats the process, except the null has no drift ($\Delta\chi_t = \epsilon_t$). The test statistic (Φ_2) is thus for the constraint that $\mu = \alpha = \beta = 0$. If Φ_2 exceeds the critical value, this is tantamount to rejecting $\mu = 0$ (given the previous test). Thus, the original null was fine, and the procedure ends by "accepting" non-stationarity on the basis of Φ_3 . But, if Φ_2 does not exceed the critical value, the appropriate null does not include a drift. However, the procedure cannot end by acceptance of driftless non-stationarity because Φ_2 is sensitive to the presence of a time trend in the alternative. Therefore, a more appropriate test may be against Alternative 2, which is:

$$\Delta\chi_t = \mu^* + \alpha^*\chi_{t-1} + \sum_{i=1}^4 \gamma_i^* \Delta\chi_{t-i} + u_t^*$$

The test statistic (Φ_1) is thus for the constraint that $\mu^* = \alpha^* = 0$ (but under a different null than Φ_3). If Φ_1 exceeds the critical value, one rejects the null hypothesis of non-stationarity and the procedure ends, Φ_1 being invariant to the presence of a non-zero mean in the alternative. If Φ_1 is below the critical value, one accepts non-stationarity after having checked, on the basis of its t-statistic, that α is not significantly less than zero in Alternative 1 (where $\mu = 0$, as the series may have a zero mean in the alternative):

$$\Delta\chi_t = \alpha'\chi_{t-1} + \sum_{i=1}^4 \gamma_i' \Delta\chi_{t-i} + u_t'$$

The column $t_{\alpha'}$ gives the t-statistics associated with the coefficient of χ_{t-1} (α') in Alternative 1. Φ_3 , Φ_2 , Φ_1 are the F-statistics associated with the joint test of the null hypotheses, respectively $(\mu, \beta, \alpha) = (\mu, 0, 0)$; $(\mu, \beta, \alpha) = (0, 0, 0)$; $(\mu^*, \alpha^*) = (0, 0)$. The critical values do not follow the usual t- and F-distributions. For 50 observations, the critical value for the t-statistic at the 5 per cent level for Alternative 1 is -1.95 (Fuller (1976), *Introduction to Statistical Time Series*, p. 373). Critical values for Φ_3 , Φ_2 and Φ_1 are 6.73, 5.13, 4.86 (Dickey and Fuller (1981), *Econometrica* 4, p. 1063).

Table B2
Cointegration tests
Dependent variable: $\ln W - \ln P - \ln E$

Right-hand side variable:	United States	Japan	Germany	France	Italy	United Kingdom	Canada
Intercept	-0.70	-1.94	-0.98	-0.21	0.34	-0.08	-0.30
T	-0.01	-0.72	--	-1.08	-0.67	-0.25	--
$\ln(P_0/P)$	0.76	--	0.48	0.15	1.00(i)	1.00(i)	0.30
Tx	--	--	0.38	1.74	--	0.68	--
POP	--	--	-0.57	--	-0.84	-0.51	-0.16
$\ln SHW$	--	-1.67	-0.82	--	--	--	--
Number of observations	59	49	53	53	57	55	45
Dickey-Fuller statistic	-3.25	-2.96	-3.46	-2.32	-2.85	-5.42	-3.28
Observations from 1980 to 1989 omitted	-2.49	-2.25	-2.08	-1.94	-2.18	-4.37*	-2.35
Observations from 1985 to 1989 omitted	-2.82	-2.63	-3.25	-2.42	-2.59	-5.53*	-3.10

An asterisk denotes that the equation passes the test at the 5% level; (i) indicates that the coefficient value was imposed.

Definition of the variables. W: business-sector wage rate; P: current consumption deflator; E: current labour efficiency; P_0 : business-sector GDP deflator; T: average effective rate of employers' contributions; Tx: average effective tax rate on personal income; POP: ratio of working-age to total population; SHW: share of women in the labour force. \ln denotes the logarithm operator.

Description of the test. Two series, x and y, integrated of order one (i.e. made stationary when differenced once), are said to be cointegrated if there exists a stationary linear combination (the cointegrating equation) of the two of them: $y = a + b \cdot x + \epsilon$ where ϵ is stationary. The concept can be extended to more than two variables.

The test is the fourth order-correction-augmented-Dickey-Fuller test. The null hypothesis of non-stationarity of the residuals ϵ_t of the cointegrating equation is rejected if the t-statistic (the Dickey-Fuller statistic in the table) associated with α is significantly less than zero in the following model:

$$\Delta \epsilon_t = \alpha \epsilon_{t-1} + \sum_{i=1}^4 \gamma_i \Delta \epsilon_{t-i} + u_t.$$

The 5% and 10% critical values above which one rejects the non-stationarity hypothesis are -3.75 and -3.36 with 3 regressors (including the constant) and -4.15 and -3.85 with 5 regressors, for 50 observations (Engle and Yoo, *Journal of Econometrics*, No. 35, 1987, Table 3). The corresponding asymptotic critical values are -4.11 and -3.83 with 3 regressors and -4.71 and -4.43 with 5 regressors (Phillips and Ouliaris, *Econometrica*, No. 1, 1990, Table IIb). Thus, if the "Dickey-Fuller" statistic is below the critical value, there is no evidence that the variables are cointegrated. OLS estimates are used. The t-statistics associated with the coefficients of the cointegrating variables are not reported as they do not follow the usual t-distribution.

ANNEX CTHE MODEL OF WAGE DETERMINATION1. Model specification

In order to ensure well-behaved long-run properties, the elasticities of the dependent variable -- the wage rate -- with respect to prices and labour efficiency have been imposed to be unity, a restriction which was easily accepted by the data. This restriction is best viewed as a specification test. If, over sample periods as long as were used for the estimations, these elasticities were found not to be one, this would not suggest that wages were under-indexed but rather that the rest of the model was misspecified. However, the relative weights of output prices, cyclically-adjusted labour efficiency and expected inflation have been freely estimated, as well as the wage sensitivity to changes in the rate of employers' social-security contributions.

On a priori grounds, the Hodrick-Prescott filtered inflation rate was used as a proxy for consumer price inflation expectations. However, if for some reason expectations are in fact perfect or static, they will be captured by the current inflation rate. Therefore, the estimated coefficient of the trend rate does not measure, strictly speaking, the room left for "expectations" in general to influence wage behaviour. With the exception of Sweden, current values of output prices were used since expected (trend) values proved to be insignificant. This may be due to the fact the gap between current and trend output price inflation does not bring much additional information as compared to the consumer price gap (for seven out of ten countries, the correlation between the two gaps exceeds 60 per cent). It is also possible that firms -- as price-setters -- are in a better position than consumers for prices to be realised close to their expectations.

Labour-market disequilibrium was proxied by the gap between the actual and trend unemployment rates. Except for Canada and the smaller countries, the logarithm of these rates was used, thereby allowing for asymmetric unemployment influences due, perhaps, to the presence of legal minimum wages and other "floor" effects. Additional variables, though of less importance, are the change in working-age population (as a deviation from the average growth rate) and the past excess of expectations over ex-post inflation, which proved to have some influence in certain countries.

Apart from the unemployment rate, all variables are expressed in growth-rate terms. In the long run, the unemployment gap is closed while wage growth and the growth of the right-hand-side variables sum to zero, so that no constant term has been included, a constraint accepted by the data.

The complete model is therefore the following:

$$[1] \quad dW = a_1 dPo + (1-a_1) [a_2 dPe + (1-a_2) dP] + a_3 dES + (1-a_3) dE + a_4 UG \\ + a_5 dT + a_6 dPS (-i) + a_7 GPOP(-j)$$

with a_1 , a_2 and a_3 being positive and a_4 , a_5 , a_6 and a_7 negative; and where W is the wage rate, P_o the output price, P and P_e current and expected consumer prices, E and ES actual and trend labour efficiency, UG the unemployment gap, T the employers' social-security contribution rate, PS price surprises, and $GPOP$ the demographic factor; d denotes the arithmetic rate of change operator.

However, if a stable long-run equilibrium level of wages exists, an equation which determines only the rate of change of wages is misspecified because the information that wages ultimately revert to this level is lost. As cointegration tests in Annex B suggest, a specification in levels may be more appropriate for the United Kingdom, and perhaps for North America and Germany. However, the influence of the unemployment gap significantly weakens when equation [1] is transformed into an error-correction model (i.e. in levels in the long run) by adding as an explanatory variable the lagged residuals from the cointegrating equations for these countries. This is due to the high correlation between the current unemployment gap and past excess real wages

relative to productivity levels as measured by the error-correction term (Chart A). Thus, the unemployment gap variable may play the role of bringing wages back to some equilibrium levels.

2. Factor contributions

Estimation results are reported in Table C1. With the exception of Canada, all coefficients are in general reasonably significant. Except for France and Spain, the D-W test is not below the critical value. Factor contributions for the 1980s are reported in Table C2. The unemployment gap term is the main transmission channel for the influence of cycles on wages, although the contribution of changes in non-cyclically adjusted productivity is not negligible for Italy. Consistent with a large body of the empirical literature, wages in Japan are more responsive to labour-market conditions than in other countries. However, given the small fluctuations of the unemployment gap, the contribution of unemployment to the variance of wage changes is no larger for Japan than for other countries.

Cyclical influences are not, however, the whole story and other factors may have been more important for some countries in certain periods. Output price inflation had a large and positive effect in Germany from 1983 to 1987, but a strongly negative effect on average in smaller European countries. Changes in employers' social-security contributions substantially depressed wages in the major continental European countries and Australia in the 1980s. Demographic shocks also seem to be relevant for understanding wage moderation in Germany and Italy. On the other hand, current and past consumer price surprises have played only a minor role, except in Spain and Sweden.

For the 1980s as a whole, none of the explanatory factors would have exerted a positive influence on the gap between current real consumption wage growth and trend labour efficiency growth, with the exception of relative output prices for Germany, employers' contributions for the United Kingdom and Sweden and demographic variables for Canada.

3. An improved equation for France

The tracking performance of the equations is shown in Chart B. While the model "fits" well for the United States and Sweden, the residuals are sometimes large for other countries (although still smaller than those from standard Phillips-curve specifications). In particular, for France the persistent tendency to overpredict from 1982 is an indication that the equation is unstable. It is substantially improved when a dummy variable accounting for wage control measures in the second half of 1982 is added and if price expectations are assumed to be purely adaptive before 1983. The new equation is the following:

$$\begin{aligned}
 [2] \quad dW = & 0.12 \, dPo + 0.88 \{ (1-D) * [0.52 \, dPa + 0.48 \, dP] + \\
 & \quad (2.0) \quad \quad \quad (7.1) \\
 & \quad D * [0.78 \, (1.14 \, dPe - 0.14 \, dPa) + 0.22 \, dP] \} - 0.21 \, (1-D) * (dPa - dP) [-1] \\
 & \quad (3.4) \quad (4.5) \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad \quad (-2.5) \\
 & + 0.95 \, dES + 0.05 \, dE - 2.91 \, UG - 0.17 \, dT - 1.9 \, D82.2 \\
 & \quad (14.4) \quad \quad \quad (-4.9) \quad \quad (-4.1) \quad (-4.5) \\
 DW = & 1.7 \quad SEE \, (\%) = 0.37 \quad adjR^2 = 0.97
 \end{aligned}$$

where dPa is a four-semester moving average of lagged inflation rates, D is a dummy variable set at unity from 1983 to 1989, $D82.2$ is a dummy variable set at unity in S2.1982. The other variables are the same as those in equation [1].

The coefficient of smoothed inflation expectations (dPe) was found to be negative, although not significant, before 1983. Therefore, this variable was omitted for the estimation until 1983. On the other hand, its coefficient is not significantly different from unity after 1983 (1.14 with a standard error of 0.25). The observation that the weight of the adaptive-expectation variable (dPa) becomes negative after 1982 may indicate that the shift towards forward-looking expectations is even more marked than measured by the two-sided smoothed inflation rate. Also, the implicit constraint that long-term indexation of wages to prices is identical before and after 1983 is not rejected by an F-test at the 10 per cent level, whether unitary indexation is imposed before 1983 ($2.33 < F[1,30]=2.88$) or the price elasticity is freely

estimated ($1.05 < F[1,29]=2.89$). On the whole, the systematic overprediction has disappeared as a dummy variable set at unity from 1983 is not significant (t-statistic = -1.0). These results are consistent with findings of a recent study using a Rational Expectation Hypothesis approach (Blanchard and Sevestre, 1989).

Table C1

Estimation results

The model: $dW = a_1 dPo + (1-a_1) [a_2 dPe + (1-a_2) dP] + a_3 dES + (1-a_3) dE + a_4 UG + a_5 dT + a_6 dPS (-i) + a_7 GPOP (-j)$

Coefficients: (t-statistic)	United States	Japan	Germany	France	Italy	United Kingdom	Canada	Australia	Spain	Sweden
a1	0.20 (1.8)	0.47 (2.6)	0.90 (9.7)	0.20 (2.3)	0.55 (2.4)	0.62 (5.8)	0.19 (1.2)	0.59 (2.2)	0.13 (1.0)	0.87 (2.6)
a2	0.74 (6.1)	0.56 (2.7)	--	0.45 (4.1)	0.67 (2.8)	0.23 (2.1)	0.41 (1.7)	--	0.90 (7.7)	0.52 (2.2)
a3	0.94 (24.4)	0.87 (6.6)	0.70 (15.8)	0.92 (10.0)	0.58 (5.2)	0.69 (9.4)	0.85 (7.5)	0.68 (4.9)	0.84 (8.9)	0.87 (7.4)
a4	-1.10 (-3.0)	-9.84 (-4.5)	-0.93 (-4.4)	-2.90 (-4.3)	-7.28 (-2.4)	-1.66 (-2.7)	-0.49 (-2.5)	-1.02 (-2.9)	-0.49 (-4.2)	-1.73 (-2.2)
a5	-0.07 (-2.9)	--	-0.20 (-3.6)	-0.23 (-4.4)	-0.26 (-3.6)	-0.10 (-2.6)	--	-0.15 (-2.1)	-0.12 (-2.4)	-0.21 (-12.3)
a6	-0.16 (-1.5)	-0.25 (-1.4)	--	-0.41 (-3.4)	-0.46 (-2.6)	-0.19 (-1.9)	-0.22 (-1.0)	--	--	-0.25 (-2.1)
a7	--	--	-1.12 (-2.9)	--	-3.07 (-3.1)	--	-0.35 (-0.8)	--	--	--
DW	1.9	1.7	2.0	1.4	1.9	2.0	1.6	1.8	0.7	1.7
SEE (%)	0.38	1.23	0.73	0.54	1.52	0.89	1.06	1.66	1.28	1.75
adj R ²	0.94	0.67	0.90	0.90	0.72	0.85	0.69	0.76	0.74	0.80
Sample period										
beginning year:	S2.1961	S2.1966	S2.1962	S1.1970	S1.1962	S2.1961	S1.1967	S1.1970	S2.1965	S2.1964
End-year: S2.1989										
F-test for the restriction of unitary price and productivity elasticities (critical value at the 5% level in parentheses)	3.49 (4.04)	1.82 (4.08)	0.56 (4.06)	2.40 (4.15)	0.14 (4.05)	2.03 (4.05)	1.03 (4.12)	1.60 (4.15)	1.74 (4.04)	1.23 (4.10)
Chow test for break from 1980 (critical value at the 5% level in parentheses)	0.29 (2.32)	1.57 (2.41)	1.15 (2.43)	1.99 (2.45)	1.30 (2.09)	1.34 (2.32)	0.45 (2.38)	2.56 (2.68)	0.43 (2.34)	2.03 (2.27)
Percentage (%) of semi-annual over-predictions 1980-89	50	55	50	55	40	45	45	50	60	55
Test for a dummy variable equal to:										
1 from 1980, 0 before	0.05 (0.5)	-0.18 (-0.6)	-0.13 (-0.8)	-0.34 (-2.5)	0.45 (1.1)	0.10 (0.5)	0.29 (0.8)	0.15 (0.3)	-0.26 (-0.7)	-0.27 (-0.6)
1 from 1985, 0 before	-0.08 (-0.6)	0.24 (0.6)	-0.25 (-1.1)	-0.10 (-0.5)	0.93 (1.7)	0.33 (1.1)	0.43 (0.9)	-0.80 (-1.5)	-0.29 (-0.6)	0.24 (0.4)

Notes:

1. W: business-sector wage rate; Po: business-sector GDP deflator; Pe: trend consumption deflator; P: current consumption deflator; ES: trend labour efficiency; E: current labour efficiency; UG: current unemployment rate less trend rate, both in logarithm except for Canada and smaller countries (the unemployment gap is one-semester lagged for Australia); T: average effective rate of employers' contributions; dPS: difference between trend and current consumption deflator change rates (business-sector GDP deflator for Sweden); GPOP: growth rate of working-age population less sample-period average growth rate.

d denotes the arithmetic rate of change operator.

2. A set of dummy variables has been included:

United States: 1.0 in 67.I and 67.II; Japan: 1.0 in 74.I and 75.I, and -1.0 in 74.II and 75.II; Germany: alternates of 1 to -1 from 71.I to 74.II; United Kingdom: 1.0 in 70.I, and 1.0 from 74.II to 75.I and -1.0 from 75.II to 77.II; Canada: 1.0 in 70.I and 1.0 from 77.II-78.II; Australia: 1 in 74.I and 74.II. See Chan-Lee et al., *OECD Economic Studies*, No. 8, Spring 1987 for the justifications of these dummy variables.

Table C2

Decomposition of the wage equations
 (Factor contributions to nominal wage changes, annual arithmetic average of half-year rates of change)

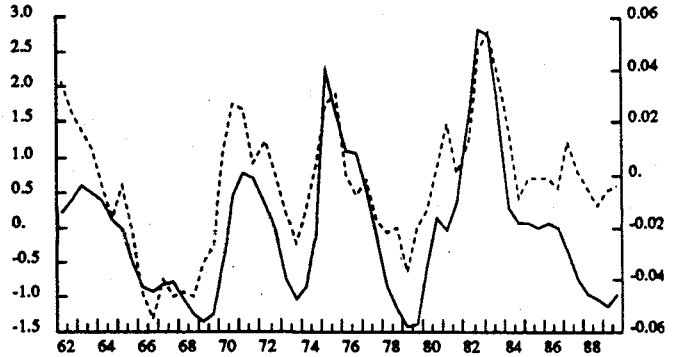
	United States				Japan				Germany				France				Italy			
	1980-1983	1984-1987	1988-1989	1990-1999	1980-1983	1984-1987	1988-1989	1990-1999	1980-1983	1984-1987	1988-1989	1990-1999	1980-1983	1984-1987	1988-1989	1990-1999	1980-1983	1984-1987	1988-1989	1990-1999
Predicted gap between current real consumption wage growth and trend labour efficiency growth	-0.7	0.1	0.5	-0.2	-0.5	-1.7	1.0	-0.7	-2.2	0.9	0.6	-0.4	-1.0	-1.0	-0.2	-0.8	-2.4	-0.9	-1.1	-1.5
of which:																				
Cyclical productivity gap (a)	0.0	0.1	0.0	0.0	-0.1	0.1	0.1	0.0	-0.5	0.2	0.4	0.0	-0.1	0.1	0.1	0.0	-0.7	0.5	0.8	0.1
Price surprises (b)	-0.2	0.0	0.2	0.0	-0.1	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.1	-0.1	0.0	-0.2	0.0	0.1	0.0
Unemployment gap	-0.3	0.0	0.4	0.0	0.4	-1.5	1.2	-0.2	0.2	-0.3	0.1	0.0	-0.1	-0.8	-0.3	-0.4	0.2	-0.3	-0.6	-0.2
Business GDP deflator (c)	-0.1	-0.1	-0.1	-0.1	-0.7	-0.3	-0.1	-0.4	-0.7	1.2	-0.3	0.2	-0.3	0.2	0.0	0.0	-0.5	0.1	-0.1	-0.2
Employers' contributions	-0.1	0.1	0.0	0.0	0.0	0.0	0.0	0.0	-0.4	-0.2	0.2	-0.2	-0.6	-0.4	0.1	-0.4	-0.9	0.1	-1.9	-0.7
Demographic changes	--	--	--	--	--	--	--	--	-0.8	-0.1	0.2	-0.3	--	--	--	--	-0.3	-1.3	0.5	-0.6
Residual (actual less predicted)	0.3	0.0	-0.3	0.1	-1.8	1.1	-0.3	-0.3	0.0	-0.1	-1.0	-0.2	-0.6	-0.4	-0.3	-0.5	0.1	0.9	1.2	0.7
Predicted gap between current real consumption wage growth and trend labour efficiency growth	-0.6	-0.2	0.6	-0.2	-0.8	0.1	1.2	-0.1	-1.4	-2.5	1.7	-1.2	-2.8	-1.8	-0.5	-1.9	-2.9	1.0	0.7	-0.6
of which:																				
Cyclical productivity gap (a)	0.0	0.3	-0.8	0.0	-0.2	0.2	-0.2	0.0	-0.4	0.3	-0.4	-0.1	-0.1	0.3	-0.1	0.1	0.0	0.1	-0.2	0.0
Price surprises (b)	0.2	0.0	-0.2	0.0	-0.3	0.2	-0.1	-0.1	--	--	--	0.0	-1.1	0.9	-0.5	-0.2	-0.2	0.0	0.6	0.1
Unemployment gap	-0.3	-0.6	0.7	-0.2	-0.2	-0.5	1.2	-0.1	0.2	-1.8	1.0	-0.4	-0.1	-0.1	0.0	-0.1	-0.3	-0.3	1.2	0.0
Business GDP deflator (c)	-0.2	-0.2	0.5	-0.1	-0.2	-0.2	0.0	-0.2	-0.1	-0.4	1.0	0.0	-1.1	-2.6	-0.3	-1.5	-1.9	0.3	-0.4	-0.7
Employers' contributions	-0.3	0.4	0.4	0.1	--	--	--	--	-1.2	-0.6	0.1	-0.7	-0.4	-0.1	0.4	-0.2	-0.5	0.9	-0.4	0.1
Demographic changes	--	--	--	--	0.1	0.3	0.3	0.3	--	--	--	--	--	--	--	--	--	--	--	--
Residual (actual less predicted)	-0.4	1.0	-0.2	0.2	-0.2	0.6	0.7	0.3	1.3	0.6	-2.6	0.2	-0.2	0.1	-1.3	-0.3	-2.1	1.1	0.1	-0.4

a) Difference between current and labour efficiency growth.
 b) Cumulated impact of past and current gaps between expected and current inflation.
 c) Relative to consumption deflator.

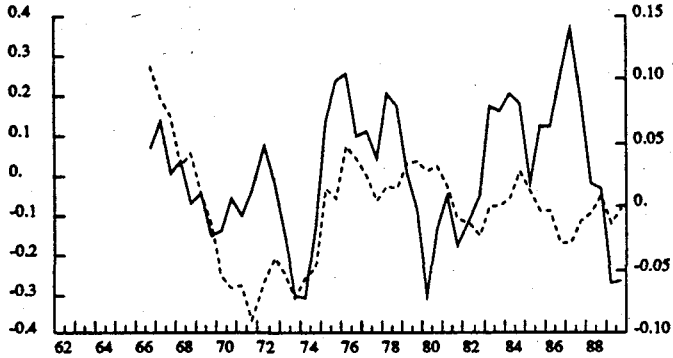
Chart A Unemployment gaps and wage deviations from 'normal' levels

— Actual less trend unemployment rate (left scale)
 - - - Lagged residuals from cointegrating equations as in Table B2 (right scale)

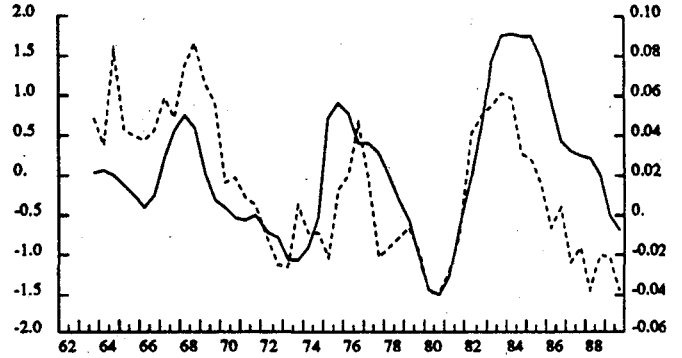
United States



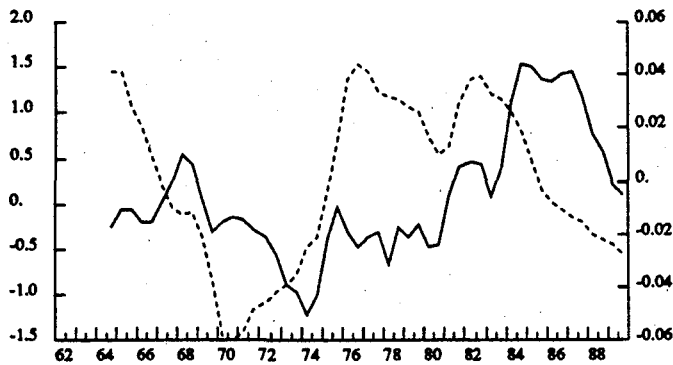
Japan



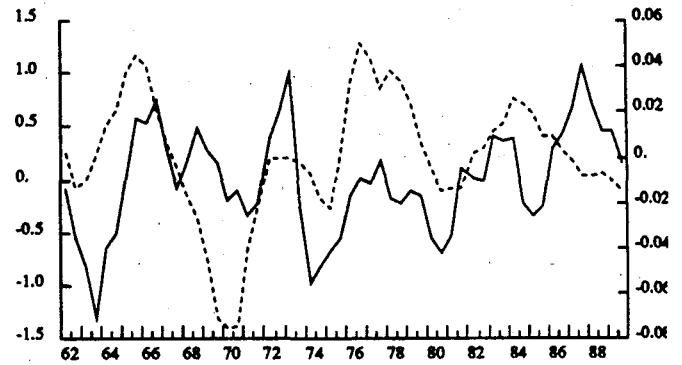
Germany



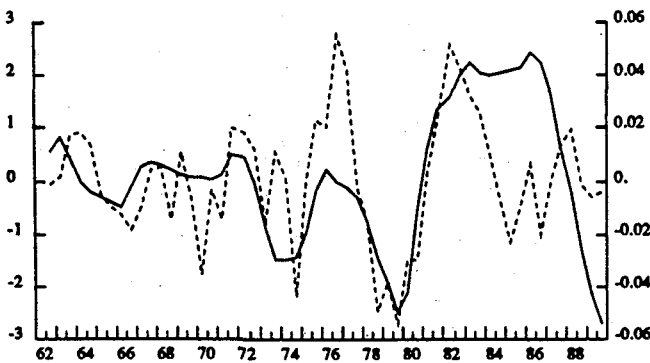
France



Italy



United Kingdom



Canada

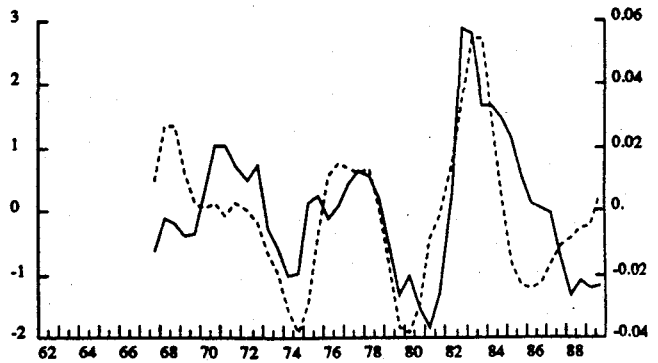
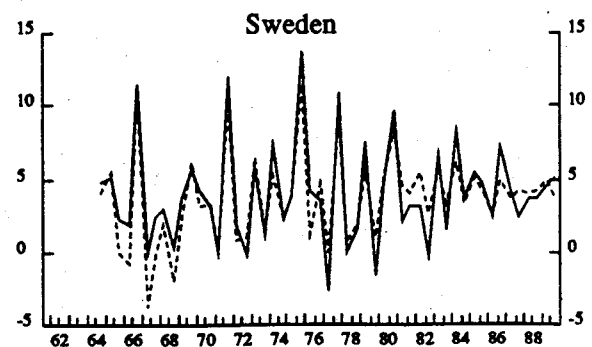
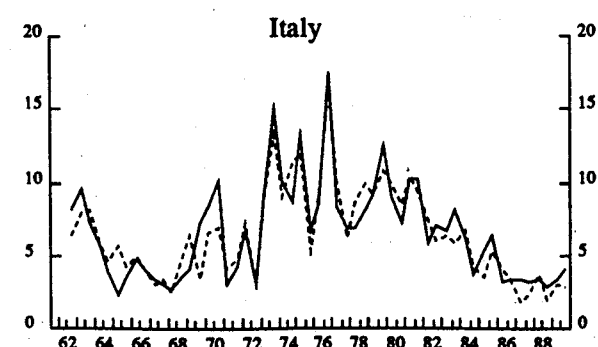
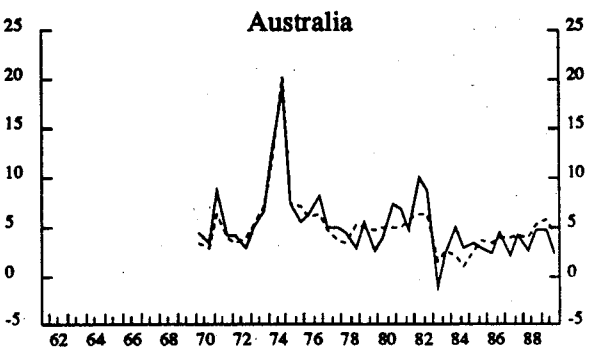
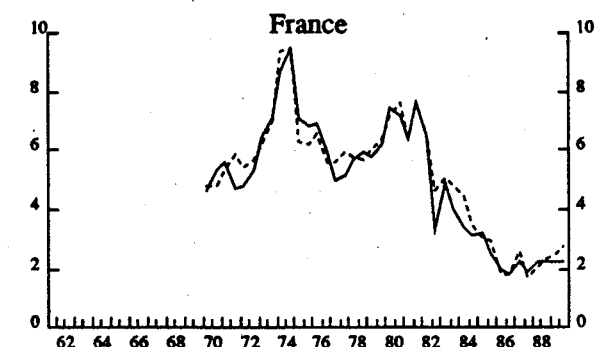
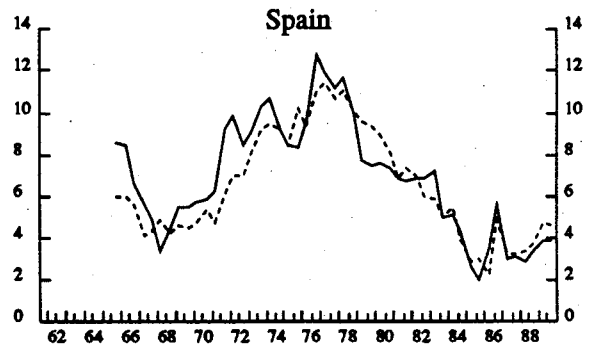
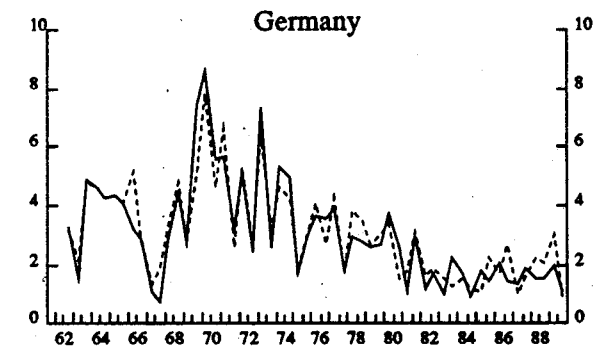
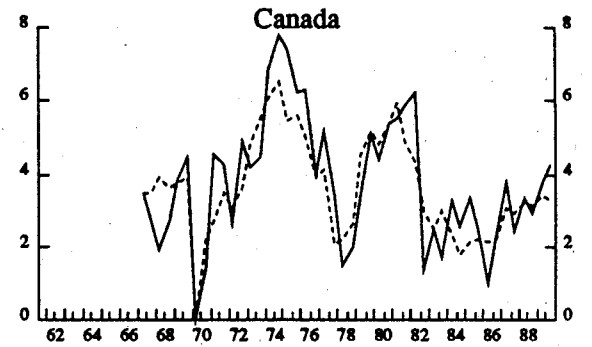
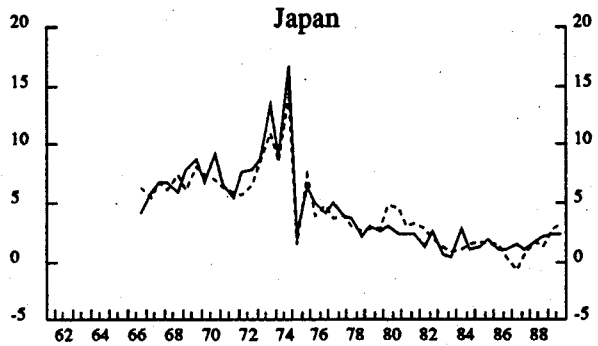
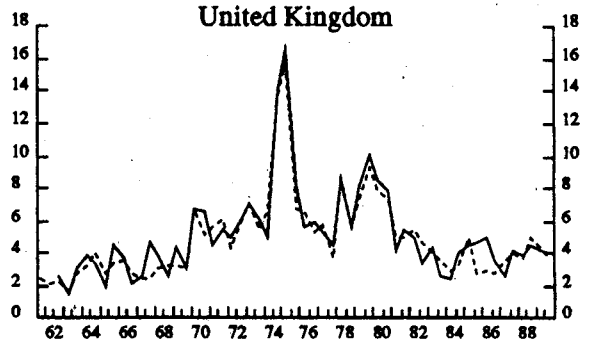
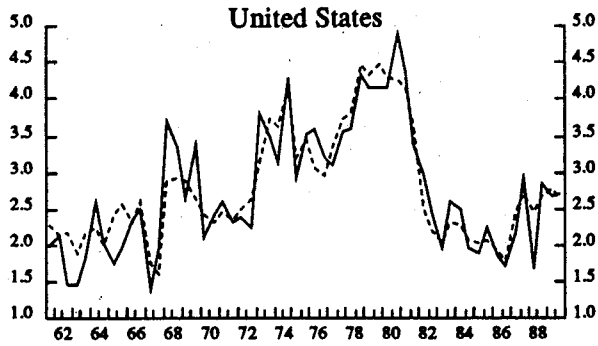


Chart B. Actual and predicted wages

(percentage changes from previous half-year)

— Actual
- - - - - Predicted



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6/11/90

ECONOMIC AND STATISTICS DEPARTMENT

WORKING PAPERS

In April 1983, the Economics and Statistics Department initiated a new series of economic studies entitled ESD Working Papers.

The following titles have been circulated:

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