Understanding the recent behaviour of inflation: an empirical study of wage and price developments in eight countries[1]

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Abstract

An important and surprising characteristic of the economies in industrialised countries in the 1990s was the extent to which prices decelerated in an environment of generally rising economic activity and tightening labour markets. The coexistence of a healthy economic environment and low inflation can appropriately be described as good news. However, economists tended to overstate underlying inflation pressures in many of these countries during most of the decade, and these prediction errors demonstrate our lack of understanding of the inflation process and raise questions about the appropriate stance of monetary policy in such an uncertain environment.

In this paper, we investigate the magnitudes and potential sources of inflation forecast errors during the 1990s in a sample of eight industrialised countries. Our analysis consists of two separate approaches. First, we examine the errors in official OECD forecasts, which we take to be representative of the mainstream of macroeconomic analysis during that time. Second, we document and analyse prediction errors in our own set of econometric specifications, which are loosely based on the Phillips curve model of the inflation process.

Our analysis of OECD forecast errors is indicative of persistent overpredictions of price inflation for most of the countries in our sample. In contrast, little bias is evident in the OECD forecasts of wage inflation. The combination of the forecast errors for wages and prices thus implies that *real* wage growth has been unexpectedly strong during the 1990s and that the major *sources* of the forecast errors are likely to be located in that part of the Phillips curve framework which models firms' prices as a mark-up on costs. More precisely, the unexpectedly slow rise in prices relative to wages could indicate that firms have benefited from favourable supply shocks or that they lost pricing power during the 1990s in that they were not able to fully pass on wage cost increases into their prices.

Our own models for consumer prices also consistently overpredict inflation in nearly every country. Nonetheless, there is little statistical evidence of parameter instability. The one exception is a decline in the intercept term in the 1990s, a finding indicative of structural change but, unfortunately, not particularly helpful in identifying the source of the change. In contrast, the parameters in our wage models appear to have changed in about half of the countries. However, no single coefficient stands out as particularly sensitive to the addition of the more recent data.

1. Introduction

An important and, to many observers, surprising characteristic of the economies in industrialised countries in the 1990s was the extent to which prices decelerated in an environment of generally rising economic activity and tightening labour markets. Measured by the private consumption deflator, the average inflation rate among OECD member countries fell from 4.6% in 1990 to just 1.2% in 1999, while the average unemployment rate for 2000 was the lowest in 10 years. Moreover, for the first time

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since 1990 actual output for the whole OECD area exceeded potential output last year. To further illustrate the degree of disinflation during the 1990s, only four industrial countries (Greece, Iceland, Ireland and Spain) had an inflation rate *higher* than 2½% in 1999, while in 1990 only three countries (Ireland, Luxembourg and the Netherlands) had an inflation rate *lower* than 2½%. Moreover, as we shall discuss in this paper, economists evidently overstated underlying inflation pressures in many of these countries during most of the decade, at least as indicated by the persistent overpredictions of price changes evident across a wide range of forecasts. While the coexistence of a healthy economic environment and low inflation can appropriately be described as good news, these prediction errors also demonstrate our lack of understanding of the inflation process and raise questions about the appropriate stance of monetary policy in such an uncertain environment.

Given that disinflation has occurred in so many countries, it might be expected that the primary reason for this favourable development would be clearly evident and pertain to a large number of - if not all - industrialised nations. Indeed, there is a general consensus as to the most likely sources of disinflation, including monetary policies aimed at price stability, declines in relative import prices, a fall in the natural rate of unemployment, increased globalisation and firms' loss of pricing power. However, estimates of the contribution of these various sources and the degree to which they apply across countries are not very precise, and thus there is less agreement as to the causes of the persistent overpredictions of inflation over this period. For example, in some countries, most notably the United States and Australia, an unanticipated increase in structural productivity growth is often cited as a key cause. However, an increase in trend productivity is an unlikely candidate for consideration in many continental European countries simply because no such improvement is evident in the statistics. Similarly, while there are signs of higher productivity growth late in the 1990s in Canada, such improvements came too late for productivity developments to be a credible source of forecast errors. In several of these countries, as well as in the United Kingdom, labour market reforms (in many cases aimed at increasing the labour intensity of output growth), inflation targeting and other structural changes are typically highlighted instead.

In this paper, we investigate the magnitudes and possible sources of inflation forecast errors during the 1990s in a sample of eight industrialised countries.[2] Our analysis consists of two separate approaches. First, we examine the errors in official OECD forecasts, which we take to be representative of the mainstream of macroeconomic analysis during that time.[3] Such an exercise helps us document the direction and magnitude of forecast errors and allows us to look at possible correlations between the surprises in wage and price inflation and surprises in other variables that are typically included in forecasters' models of the inflation process. We then turn to our own set of econometric specifications, which are loosely based on the Phillips curve model of the inflation process. We again document the size of prediction errors made by these models, but we also use them to test for possible structural change, either in the NAIRU (non-accelerating inflation rate of unemployment) or in the responsiveness of wage and price inflation to changes in the various determinants of inflation included in the models.

To preview the results, our analysis of OECD forecast errors is indicative of persistent overpredictions of price inflation for many - although not all - of the countries included in our sample; especially large forecast biases are evident for the United States, Canada and Australia. In contrast to the results for prices, little bias is evident in the OECD forecasts of wage inflation, with the notable exception of Japan. For most countries, the combination of the forecast errors for wages and prices implies that *real* wage growth was unexpectedly strong during the 1990s and that the major *sources* of the forecast errors are likely to be located in that part of the Phillips curve framework which models firms' prices as a mark-up on costs. Thus, the unexpectedly slow rise in prices relative to wages could indicate that firms have benefited from favourable supply shocks or that they lost pricing power during the 1990s in that they were not able to fully pass on wage cost increases into their prices. In contrast, the part of the Phillips curve framework that models wages as a function of labour market slack and the expected rate of inflation seems to contain fewer systematic errors.

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2 The countries included in our analysis were selected because of the participation of their central banks in the workshop.

3 See Batchelor (2000), who analyses the accuracy of forecasts for the Group of Seven countries made by the OECD, the IMF and the consensus of private sector economists published by Consensus Economics. While the private sector forecasts appear to be more accurate than those made by the OECD and the IMF, all three groups have tended to overpredict inflation in the 1990s.

We also compared the forecast errors for wage and price inflation with the levels and forecast errors for some of the standard determinants of inflation. In particular, we first analysed the correlation matrix of these forecast errors, but found no dominant explanation for the OECD's overly pessimistic view of inflation pressures in the 1990s. Larger than expected declines in relative import prices seem to be a relevant explanation for some countries. However, a number of countries for which the bias in the OECD inflation forecasts was small also benefited from lower import prices. For other countries, an acceleration in productivity appears to be an important factor, while in still others, errors in forecasts of the GDP gap and/or the unemployment rate appear to play a role. We also regressed the forecast errors directly on the exogenous determinants of inflation to test for structural changes. Again, the results varied across countries. Although there was evidence of some structural change in most cases, the specific nature of the parameter changes differed considerably, ranging from a step-down in the average rate of inflation to changes in the responsiveness of wages or prices to supply shocks or cyclical influences.

Turning to our own specifications, the estimated models for consumer prices consistently overpredict inflation in nearly every country. Nonetheless, there is little statistical evidence of parameter instability in these models. The one exception is a decline in the intercept term in the 1990s, a finding indicative of structural change but, unfortunately, not particularly helpful in identifying the source of the change. In contrast to the price models, the cumulative forecast errors from our models of wage inflation are negative in some countries but positive in others. Moreover, the parameters in about half of the countries appear to have changed during the 1990s, although no single coefficient stands out as particularly sensitive to the addition of the more recent data.

2. Alternative explanations of inflation developments in the 1990s

Recent characterisations of macroeconomic conditions have been influenced to a significant degree by developments in the United States, where inflation and unemployment fell simultaneously over much of the 1990s. Accordingly, many of the hypotheses advanced to explain this favourable economic performance emphasise the US experience and underweight important developments elsewhere in the industrialised world. That said, the following brief review also focuses mainly on explanations put forward with regard to the US economy, although we attempt to include, where appropriate, other hypotheses that seem more relevant elsewhere.

In general, the recent literature on this issue can be loosely classified into three broad categories.[4] The first includes arguments that the inflation process has irrevocably changed; on this view, the models that economists have traditionally used to forecast inflation are no longer applicable. At the other end of the spectrum are claims that recent inflation developments are easily explained in the context of standard inflation models as long as various one-time supply shocks are accounted for. The third category comprises those in the middle who argue that the underlying model of inflation still holds, but that some of the key parameters may have changed. Because this group is very large, it is useful to divide it into several subcategories, depending on the parameters of interest and the nature of the changes. For instance, some parameter shifts may at first glance seem permanent, but are in fact the manifestation of very long adjustment lags to either transitory or permanent changes in other variables.

Returning to the first category and the argument that the inflation process has irrevocably changed, such views often appear in the popular press with little explanation. But more serious proposals along these lines seem to have their roots in the way that information technology has altered traditional relationships between economic agents. For example, some proponents of this view cite DeLong and Froomkin (1999), who argue that recent technological advances in electronic commerce and data communications technology are leading to changes in how competitive markets function that may have important implications for firms' price setting behaviour.

[4] Although most of the players and the arguments are different, this debate in many respects resembles that of the late 1970s. At that time it was also postulated that because of structural changes (notably with respect to the modelling of expectations) the inflation process had fundamentally changed and the Phillips curve had broken down. Others, however, argued that as long as one-time supply shocks were taken into account, the Phillips curve was still alive and well.

Regarding claims that recent inflation developments are easily explained in the context of the standard models of inflation, Rich and Rissmiller (2000) argue that a traditional Phillips curve model tracks US inflation quite well during the 1990s and that a significant decline in import prices in that period is the main source of the deceleration in prices. However, according to others, the decline in relative import prices does not fully explain the decline in US inflation. For instance, Brayton et al (1999) find that an equation including changes in relative import prices still tends to overpredict inflation, and the same is true for our own estimates presented in Section 4.[5] Similarly, Gordon (1998) suggests that, in addition to import prices, a substantial portion of the deceleration in prices is due to a combination of unusually sharp declines in computer prices, extremely modest growth in the cost of medical care, and a reduction in measured inflation relative to true inflation. Gordon further notes that the decline in US inflation is not surprising given changes in capacity utilisation over the period. Leaving aside ambiguities about the precise size of the shocks, one implication of these explanations is that the shocks generating the favourable inflation outcomes may well be transitory and subject to reversal. This contrasts sharply with the polar "new economy" view that low inflation represents a more fundamental change in wage and price setting behaviour.

It is helpful to organise a review of the alternative explanations of inflation developments in the third category around the standard reduced-form equations for price and wage inflation:

$$\Delta w_t = \mu + \chi \sum_i (U - U^*)_{t-i} + \sum \delta_j \Delta w_{t-j} + \phi z_t + \eta \tag{1}$$

$$\Delta pc_t = \alpha + \beta \sum_i Gap_{t-i} + \sum_j \varphi_j \Delta pc_{t-j} + \gamma z'_t + \varepsilon_t \tag{2}$$

where Δpc and Δw denote price and wage inflation respectively, Gap denotes the output gap, U-U* is the unemployment gap (with U* a measure of the NAIRU), *z* and *z*' are supply side disturbances, and *i* and *j* are lags in years.

At its simplest level, the third category would include the argument that the NAIRU has declined in recent years.[6] Such a decline is variously attributed to a number of sources, including the impact of technological change on productivity growth. While the consensus view is that higher productivity growth has helped to keep inflation low and thus prolonged the expansion in several countries, it is debatable whether its impact on the NAIRU is permanent or just the result of very long adjustment lags. As pointed out by Blanchard and Katz (1997), productivity growth is neutral with respect to the NAIRU in the long run if wage aspirations and reservation wages adjust to the change in productivity growth. However, experience suggests that this adjustment is rather slow, so that if prices are set as a mark-up on unit labour costs, price inflation will be lower during a period of adjustment, with the slowdown most likely to be reflected in a lower intercept term (α).[7] Alternatively, since the output level associated with stable inflation temporarily increases, the effect of higher productivity growth may be seen as a decline in the sensitivity of price inflation to the output gap (β).

Another popular explanation for the recent good inflation performance involves increased globalisation and an associated rise in competition in both national and international markets. To the extent that these changes are permanent, they might affect most of the parameters in the above equations. For

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5 When using the impact coefficients from our own equations to estimate the contributions of import shocks (incorporating the effects of lower foreign prices as well as movements in exchange rates) to domestic inflation, we find particularly large effects for the United States and the United Kingdom, most notably during the latter half of the 1990s. For the other six countries, the decline in price inflation during the late 1990s mainly appears to reflect domestic factors.

6 Although we have placed Gordon (1998) in the previous category, he also finds some decline in the NAIRU even after making adjustments for the five supply shocks he considers. Changes in labour markets, such as the increased use of temporary workers and a higher rate of incarceration of previously unemployed individuals (Katz and Krueger, 1999) have also been mentioned as sources of changes in the NAIRU in the United States, as has hysteresis (Ball, 1999).

7 The adjustment lag can perhaps be partly attributed to the difficulty of identifying the nature and size of changes in productivity growth. For instance, efficiency gains obtained through job cuts or the elimination of non-profitable firms and activities will not have the same effect on real wage claims as technological progress. In addition, since a change in aggregate productivity gains often results from staggered shifts in the level of output per hour in individual firms or sectors, an alternative way of capturing the adjustment process would be to augment the wage equation with an error correction term written as $\theta \log((W/P)/Q)_{t-1}$, where Q denotes the level of productivity and θ is negative. Price and wage equations with error correction terms are discussed at greater length later in this section as well as in Section 4.

instance, the sensitivity of domestic prices and wages to changes in relative import prices (γ and ϕ) would be likely to increase. Moreover, a tendency for increased competition to curb firms' and wage earners' real income aspirations might be reflected in a reduction of the intercept terms α and μ or, perhaps, in lower sensitivities to domestic market conditions (β and χ).

Structural policies and an associated greater influence of market forces comprise another set of explanations for the disinflation of the 1990s. In theory, such policies can improve the growth-inflation trade-off through several interdependent channels (Dicks, 2000), but in practice, this explanation is probably more relevant to the European countries in our sample than to the United States (Siebert, 1997).[8] As labour markets are freed up, one would expect a greater influence of market forces to lower the NAIRU and possibly increase the sensitivity to labour market slack (χ). Similarly, in less regulated product markets the output level associated with stable inflation would probably increase. Moreover, the demand curves facing each firm are likely to be more elastic, encouraging or forcing firms to lower their mark-ups. In practice, however, the effects of structural policies overlap with many of the hypotheses considered above and they are also difficult to quantify. In addition, since the short-run effects of deregulation and other structural measures on the real economy are frequently negative, the time horizon over which the policies and their effects are analysed is of particular importance.

As highlighted in several other contributions, the coefficients in the typical forecasting equations may, in part, depend on the level of inflation, though the sign of this influence is ambiguous. According to Ball et al (1988), Lucas (1972) and Hutchison and Walsh (1998), nominal rigidities tend to increase when inflation declines. However, more recent proponents of this view have argued that, in a low-inflation regime, the sum of the coefficients on the lagged inflation terms ($\Sigma\varphi_j$ and $\Sigma\delta_j$) tends to be less than unity. This would imply that forecasts should be based on the "old-fashioned" downward-sloping Phillips curve rather than on an equation with a long-run vertical slope. For example, Taylor (2000) argues that the erosion in firms' pricing power during the late 1990s mainly resulted from a decline in the persistence of inflation in a low-inflation regime. This lower persistence is likely to have raised the elasticity of the demand curve facing firms and forced or encouraged them to reduce mark-ups. Moreover, in a less persistent inflation environment, firms may reduce the pass-through of costs into their own prices, which would be reflected in a smaller coefficient on past inflation.[9] In a low-inflation economy, agents may also view exchange rate changes as mainly transitory, with the result that domestic prices become less sensitive to exchange rate shocks. Indeed, several event studies (the United Kingdom and Sweden following the ERM crisis in 1992, Australia in the aftermath of the Asian crisis and Brazil in early 1999) suggest that the pass-through of exchange rate changes may have fallen in the 1990s.

Others have focused on the process determining expectations of inflation or on the way in which agents react to such expectations. For instance, Roberts (1997) finds that the expectation formation process is not entirely rational. Akerlof et al (2000) go one step further by arguing that workers and firms do not fully utilise information about inflation in determining wages and prices during periods of persistently low inflation. Their hypothesis is supported by empirical estimates and results in dynamics that are very similar to those highlighted by Taylor and thus deviate importantly from the traditional accelerationist models of the inflation process.[10] The regime dependence of the coefficients is also evident in Brainard and Perry (1999), who specify a model of US inflation that allows for time variation

[8] In the United States, many regulations were removed during the 1970s and 1980s, while UK product and labour markets underwent major deregulation during the 1980s. In contrast, markets in most continental European countries were still subject to various regulatory constraints and rigidities through much of the 1990s. However, there are signs that some of these constraints are being removed or have lost their effectiveness. For instance, several countries have eased working hour regulations, and companies are increasingly taking advantage of these measures to introduce more flexible labour contracts. Thus, the proportion of both temporary and part-time jobs in the EU countries increased sharply during the 1990s, with temporary job contracts particularly prominent in Spain and part-time jobs in the Netherlands. Moreover, the output growth required to keep the rate of unemployment stable in the euro area seems to have declined significantly during the 1990s, though the precise source of this change is unclear and probably differs across member countries (Schnabel, 2000).

[9] In Taylor's model, a profit-maximising firm sets its optimal price as $x_t = .125\ \Sigma(Et\ c_{t+i} + E_t\ p_{t+i} + E_t\ \varepsilon_{t+i})$, with c denoting marginal costs, p prices of other firms, ε a random error, E expectations and i = 0…3. If c_t (or p_t) follows a simple first-order regression ($c_t = \rho c_{t-1} + \mu_t$), the pass-through or mark-up coefficient would be $.125\ (1 + \rho + \rho^2 + \rho^3 + \rho^4 + \ldots)$. Consequently, less persistence (ie a lower ρ) will lead to a lower pass-through. See Taylor, op cit, pp 10-11.

[10] The model has the further implication that the unemployment rate associated with the optimal level of inflation is lower than typical measures of the NAIRU.

in several of the key parameters and find that the estimated coefficient on lagged inflation declines substantially in the 1990s.[11]

Finally, Brayton et al (1999) focus on changes in the price mark-up in explaining US inflation during the 1990s. By augmenting a traditional Phillips curve with an error correction term, measured as the inverse of real unit labour costs, they find evidence of an unusually high price mark-up during the 1990s, and argue that this factor, rather than a falling NAIRU, explains much of the disinflation over the last decade. Similarly, Gali et al (2000) rely on a mark-up model in analysing inflation in the euro area. While their model successfully explains price inflation, it does so by making prices conditional on labour costs rather than on economic activity, whereas in the model by Brayton et al both determinants are present. As noted by Roberts (1999), however, this splitting of the Phillips curve into a price and a wage component entails a risk that the former is well explained while developments in the latter (and potential sources of prediction errors) may not be given sufficient attention.

All in all, the range of potential explanations reviewed in this section highlights the uncertainty faced by policymakers in assessing current inflation risks and provides the motivation for the remainder of this paper. However, it should be stressed that while trying to distinguish between various explanations is intuitively appealing, such distinctions are often difficult to verify empirically. This is particularly true when considering the different implications of transitory influences and permanent changes and it is especially relevant to the 1990s, when a number of favourable supply shocks occurred more or less simultaneously with other, more permanent changes. While a model would ideally allow for the influence of both structural shifts and transitory supply shocks, this is rarely possible in practice and is particularly difficult when the sample period is very short. Nonetheless, we will attempt to highlight patterns that seem to us relevant to the question of whether the behaviour of inflation changed in the 1990s in the eight countries under consideration and, when it did, of the extent to which such a change was the result of structural developments as opposed to temporary and potentially reversible supply shocks.

3. Forecast errors in the 1990s

a. The OECD forecasts

In this section, we analyse the size and sign of inflation forecast errors, using forecasts of wage and price developments made by the OECD. These forecasts are prepared twice each year and presented in the semiannual *Economic Outlook*. For the purposes of this study, we take the forecasts for the following year as of June of the current year and compare them to the actual outcome as currently published.[12] Throughout this section, we measure forecast errors as *forecast* less *outcome* so that a positive error indicates that inflation has been overpredicted.

We should emphasise that we have not chosen this source with the intention of criticising the OECD forecasting procedures and record. On the contrary, the OECD forecasts offer several advantages over possible alternatives. For example:

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[11] When they apply a Kalman filter procedure that uses all observations in the sample period (1965-98), both the intercepts and the coefficients with respect to the unemployment gap are remarkably stable. In contrast, the sum of coefficients on lagged inflation varies significantly over the sample period, in the case of wage inflation rising from about 0.35 in the mid-1960s to around 0.55 in 1980 and then dropping to just above 0.30 in the late 1990s. For price inflation the change is equally pronounced, with the influence of lagged inflation increasing from just below 0.60 in the 1960s to 0.80 in 1980 and then dropping to just above 0.40 by the end of the sample period.

[12] To be more precise, the forecasts for year t refer to those made by the OECD in June of year t-1, while the outcomes for year t are those reported in the Economic Outlook as of June 2000. This differs from other evaluations of forecasts, which typically measure outcomes as of the data reported at time t. Given that we are interested in the sources of the forecast errors rather than the quality of the forecasts, we view our definition - which takes the revised data as the best measure of the true outcome - as the most relevant comparator to the forecast values. However, the results in this section were qualitatively similar when the forecast errors were defined using the outcomes as of year t.

- They provide a consistent set of time series with few breaks and virtually no changes in the dates of the forecasts;
- They include forecasts of a number of additional variables that may be relevant when searching for possible sources of errors;
- Forecast errors made by the OECD are generally of the same magnitude as those made by other international institutions, for instance the IMF and the European Commission;
- Given the semiannual meetings of forecasters held at the OECD, the forecasts presented in the *Economic Outlook* are fairly representative of official forecasts for individual countries.[13]

b. Properties of the inflation forecast errors

Tables 2 and 3 present the forecast errors for the period 1991-99 for changes in the consumption deflator (Δpc) and compensation per employee in the business sector (Δw) respectively. The tables are organised to show four standard measures used in forecast evaluation: accuracy, bias, efficiency and serial correlation. However, as we are primarily interested in the extent to which conventional forecasts overpredicted inflation during the 1990s, we shall focus on the bias and efficiency statistics.

Turning first to the results for price inflation (Table 2), the mean errors indicate that consumer price inflation was overpredicted in all of the eight countries we analyse except for Spain and the United Kingdom. Especially large and statistically significant mean errors are evident for Australia, the United States and Canada. In each of these countries, the U^M statistic indicates that 50% or more of the total mean squared error is due to the bias in the forecast. Moreover, as indicated by the % > 0 line, the forecasts were too high for at least eight of the nine years considered. In spite of smaller average prediction errors, the forecasts for several other countries also exhibit a tendency to overpredict inflation. In particular, the forecasts for Japan, Sweden and Switzerland were too high more than 75% of the time, possibly indicating that systematic errors in the forecasting procedure were also present for these countries.

The existence of bias in the inflation forecasts for some countries also raises doubts about the efficiency of these forecasts. In particular, an optimal forecasting procedure would yield forecast errors with specific, well defined properties that can be derived through an optimisation problem. This can be illustrated by assuming a linear relationship between the forecasts (y^f) and the outcomes (y):

$$y_t = \alpha + \beta y_t^f + \varepsilon \tag{3}$$

and rearranging this with the forecast error as the left-hand variable:

$$y_t^f - y_t = -\alpha + (1-\beta) y_t^f + \varepsilon \tag{4}$$

From this characterisation, the optimality conditions are seen to be consistent with the usual statistical criterion for minimising the sum of squared errors and can thus be stated as:

$\alpha : E(y_t^f - y_t) = 0$ and

$1-\beta : E((y_t^f - y_t) y_t^f) = 0$

where *E* denotes expected values. The first condition requires the average forecast errors to be zero; in other words optimal forecasts should be unbiased. According to the second condition, the forecasts should also take account of all existing information. A necessary condition for this to be satisfied is that (1-β) = 0. In addition, current forecast errors should not repeat past errors, so a further requirement for

[13] As an illustration, Table 1 shows forecasts and outcomes for the US Federal Open Market Committee (FOMC) and the OECD respectively. While the average forecast error for the OECD (0.8) is almost twice as high as that of the FOMC (0.45) both institutions tended to overpredict inflation in the 1990s.

efficiency is that the forecast errors are not serially correlated; we test for this using the Ljung-Box Q-statistic.[14]

Given that unbiasedness is a necessary condition for efficiency, it is not surprising that the forecasts for the three countries for which the mean error is positive and statistically significant are also not efficient. However, despite the small average errors, it also appears that the OECD forecasts for Switzerland are inefficient. In contrast, there is no evidence of either inefficiency or serial correlation in the OECD forecasts of inflation for the other four countries.

Some additional insights can be gleaned from the pattern of forecast errors shown in Chart 1. For example, while the average forecast errors for Sweden and Switzerland are rather small, each of these countries shows sizeable negative errors early in the 1990s, followed by a string of positive errors over the remainder of the decade. Similarly, for Japan, positive errors for much of the decade are largely offset by a large negative error in 1997, presumably because the introduction of a sales tax pushed up inflation by more than predicted by the OECD. Even for Spain, the underprediction of price inflation is mostly confined to the early part of the decade; in more recent years, the OECD has tended to overpredict inflation, although the errors are relatively small. With regard to the United States and Australia, the OECD forecasts show no obvious tendency to improve or worsen, while the errors for Canada, albeit still generally positive, have mainly fallen over time.

Turning to the OECD forecasts of wage inflation (Table 3), the largest average errors are observed for Japan and Australia, which are also the only countries for which the mean is statistically different from zero. Indeed, according to the measures we report, there is little evidence of a significant bias in the wage forecasts in any of the other six countries. Nonetheless, although the average forecast errors for the United States and Spain are relatively small, the pattern of errors in both countries fails the efficiency test. Among the other countries, forecast errors for Canada are predominantly positive, albeit small, while the forecasts for the United Kingdom, Sweden, and Switzerland satisfy all of the standard optimality conditions.

As shown in Chart 2, the forecast errors for compensation growth exhibit some interesting patterns. In the United States and Switzerland, for example, the OECD tended to overpredict compensation growth during the first half of the 1990s, but more recent forecasts have shown no upward bias despite the persistent positive errors in price forecasts. A similar pattern of positive errors in the first half of the 1990s is evident for Canada and the United Kingdom, although in the latter case the OECD has consistently underpredicted wage growth in recent years. Among the other countries, the forecasts for both Japan and Australia are consistently too high throughout the decade, while a tendency to underpredict wage growth in Spain in the early part of the 1990s gave way to persistent overpredictions beginning in 1994.

For the majority of countries in our sample, there seems to have been a greater tendency by the OECD to overpredict price inflation than wage inflation, a pattern that is especially apparent when the early 1990s are excluded from the time period we consider. Taken alone, and assuming that the OECD forecasts are representative of mainstream expectations, our analysis of the OECD forecast errors supports some explanations for the favourable performance of inflation in the 1990s more than others. In particular, in the absence of clear evidence of persistent positive errors in the wage forecasts, it is difficult to argue that there have been NAIRU declines associated with labour market changes that were larger than those expected by the OECD forecasters. In contrast, the upward biases in the forecasts for price inflation in many countries arguably point to a role for unanticipated supply shocks of the type emphasised by Gordon (1998) and Rich and Rissmiller (2000) or for globalisation and associated pressures on the mark-up, as suggested by Brayton et al (1999) and Taylor (2000). Of course, the absence of a consistent pattern of forecast errors across countries raises some questions about these interpretations, although the primary exceptions (Japan and Spain) appear to have been unduly influenced by atypical domestic developments.

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[14] This concept of efficiency is a relatively weak one in that it excludes a test of whether the information contained in other variables has been used. In addition, as the regression estimates of α and $(1-\beta)$ are likely to be correlated, a joint test of their significance is required. Interested readers are referred to Wallis (1989) and Barrionuevo (1992) for additional information.

c. A decomposition of the forecast errors

To investigate whether unexpected, but now identifiable, supply shocks were an important contributor to the surprisingly low inflation rates of the 1990s or whether the forecast errors tend to be more closely associated with changes in the parameters of the underlying structural model, we decomposed the OECD forecast errors into two separate components. In particular, letting the subscripts *f* and *a* denote the forecast and actual values respectively, we assume that the process used by the OECD to generate the forecast value of inflation (*y*) can be represented as:

$$y_f = \alpha_f + X_f \beta_f \tag{5}$$

where *X* represents the vector of exogenous variables used in the forecasting procedure. Similarly, the process generating the outcomes is assumed to be represented as:

$$y_a = \alpha_a + X_a \beta_a \tag{6}$$

The forecast error is the difference between these two equations:

$$y_f - y_a = (\alpha_f - \alpha_a) + X_f \beta_f - X_a \beta_a \tag{7}$$

which can be rewritten as:

$$y_f - y_a = (\alpha_f - \alpha_a) + X_a(\beta_f - \beta_a) + (X_f - X_a)\beta_f \tag{8}$$

The first two terms on the right-hand side of this equation provide an estimate of how the parameters generating the outcomes differ from those generating the OECD forecasts and can be thought of as an approximation to structural changes that have led to errors in the inflation forecasts. The last term provides an estimate of the extent to which errors in OECD forecasts of the exogenous variables contributed to the forecast errors.[15]

Can forecast errors in other variables explain the overprediction of inflation?

Because we have only nine observations with which to work, we examine the two major pieces of this decomposition separately. We consider first the potential effects of errors made by the OECD in its forecasts of variables that could be considered inputs into standard models of wage and price inflation. For instance, if import price inflation is overpredicted and import prices enter the price equation with a positive coefficient, this might explain an overprediction of consumer prices as well. Similarly, on the assumption that prices are set as a mark-up on unit labour costs, an underprediction of labour productivity growth would contribute to an overprediction of price inflation. In contrast, underpredicting labour productivity growth might be expected to cause an underprediction of wage inflation if workers alter their wage aspirations in response to rising productivity. Forecast errors for the degree of slack could also influence inflation forecasts. For instance, if the rate of unemployment were overpredicted, either wage or price inflation might be expected to be underpredicted, giving rise to a *negative* partial correlation between the prediction errors. Conversely, if the output gap turns out to be larger (ie more negative) than predicted, the forecast error on inflation would also be positive, generating a *positive* partial correlation between the errors.

In Table 4 we present the average values and forecast errors over the 1991-99 period for the four exogenous variables considered. The strongest evidence for the supply shock hypothesis comes from the fact that the OECD, on average, overpredicted import price inflation in every country during the 1990s. Similarly, average productivity growth was underpredicted (negative forecast errors) for the United States, Australia and Sweden; in contrast, there are large positive and significant forecast errors for productivity growth in Japan and Switzerland. Finally, in all countries, the OECD

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15 This procedure obviously makes strong assumptions about the relationship between the OECD forecasting procedures and the data-generating process. In particular, we are implicitly assuming that the forecasts and outcomes are both generated by the same set of variables so that there are no omitted variables. In addition, characterising differences in the parameter values as structural changes implies that the OECD estimates of the parameters at the time the forecasts are prepared represent the true values for the period prior to the forecast. Moreover, as Andersen (1997) notes, it is generally not possible to definitively identify the sources of forecast errors in the framework we are using. For these reasons, the results in this and the following subsection are intended to be suggestive rather than explicit tests of particular hypotheses.

underpredicted the unemployment rate (negative forecast errors), although only five of these countries show a commensurate positive forecast error in the degree of output slack.

As a more formal method of assessing the importance of these factors, Table 5 presents bivariate regressions of the forecast errors in price and wage inflation on the forecast errors for the four exogenous variables, along with regressions of the errors in the price forecasts on the errors in the wage forecasts and vice versa. Interestingly, there are only two cases (Canada and Australia) where forecast errors for price and wage inflation are significantly and positively correlated. In contrast, a positive correlation between forecast errors for import prices and consumer prices is evident for most countries. Indeed, given the size of the change in import prices and the associated forecast errors shown in Table 4, the surprising weakness in import prices appears to have been an important source of the overpredictions of price inflation.

The results also support the hypothesis that unexpectedly strong productivity growth contributed to the overprediction of inflation in some countries. In particular, in both the United States and Australia, the OECD underpredicted productivity growth in the 1990s, and in both countries, this underprediction appears to be associated with the positive forecast errors for price inflation. The correlation between the forecast errors for productivity and price inflation is also negative and statistically significant in Japan, although in this case, an overprediction of productivity is systematically related to an underprediction of price inflation.

Turning to the regressions of forecast errors of wage and price inflation on forecast errors of unemployment, we find three countries (the United States, Japan and Australia) with the expected negative correlations but also two cases (the United Kingdom and Spain) with a positive correlation. For the gap measure, positive and significant coefficients were found for the United States and Japan, which are consistent with the results for unemployment. However, we again find coefficients of opposite sign for the United Kingdom and Spain. In addition, the results for the United States are consistent with the traditional view that wages are more sensitive to utilisation rates than prices.

To what extent do these correlations of forecast errors "explain" the overpredictions of inflation during the 1990s? Focusing on those countries for which Tables 2 and 3 suggested a systematic overprediction, we would summarise the results in Tables 4 and 5 as follows:

- For the *United States*, a faster rate of decline in import prices than expected by the OECD and a faster rate of productivity growth both appear to have been important sources of the overprediction of price inflation. However, there are two caveats we would add to this interpretation. First, much of the rapid decline in import prices over the period occurred in imported investment goods and thus is only indirectly linked to consumer price inflation. This suggests that some of the correlation between falling import prices and consumer price disinflation may be coincidental. Second, the results indicate that the underprediction of productivity growth contributed both to the overprediction of price inflation and the underprediction of wage inflation in the 1990s. In other words, while external shocks are important to the US inflation story, it appears to be productivity shocks rather than import prices that have been the driving force.
- For *Australia*, a very similar picture emerges, as productivity growth in the 1990s seems to have been significantly underpredicted and this was transmitted into an overprediction of consumer prices. It also appears that the interaction between price and wage inflation is an important part of the disinflation process and of the inaccuracy of the forecasts, as the bivariate correlation of the forecast errors for wages and prices is sizeable and statistically significant.
- For *Japan*, higher than expected import price inflation contributes the most to the overprediction of price inflation. The forecast errors for productivity growth are significant but, in contrast to the two countries discussed above, productivity growth has been overpredicted. This helped to reduce the overprediction of price inflation in that country, but also contributed to the overprediction of wage inflation. In addition, the OECD tended to underpredict the unemployment rate during the 1990s, which also added to the positive forecast errors for wages.
- For the *United Kingdom*, the results are difficult to relate to specific hypotheses discussed in Section 2. In particular, price inflation was underpredicted slightly despite a tendency by the OECD to overpredict the degree of slack during the 1990s. Some overprediction of productivity, however, may be one source of the negative forecast errors for inflation.

- For *Canada*, a more negative than expected output gap and a positive correlation between price and wage inflation appear to be the primary sources of the forecast error for price inflation.
- For *Spain*, we find a significant positive correlation between forecast errors for productivity growth and wage inflation. This could potentially explain the overprediction of the latter, even though the average forecast error for productivity growth is very small. On the other hand, the underprediction of slack in both labour and product markets, coupled with the correlations shown in Table 5, implies a tendency to underpredict inflation.
- For *Sweden*, part of the overprediction of price inflation can be explained by an overprediction of import prices. At the same time, while productivity growth appears to be significantly underpredicted in Table 4, the corresponding bivariate regression coefficients are small and insignificant. Similarly, the rise in the estimated degree of slack (as measured by the output gap) and the associated forecast error do not seem to have contributed to the forecast errors for inflation.
- For *Switzerland*, a steeper fall in import prices than predicted by the OECD contributed to an overprediction of price inflation, as did the overprediction of wage inflation combined with relatively strong price-wage interaction effects. With regard to wage inflation, part of the positive bias appears to be related to an underprediction of the unemployment rate.

Evidence of structural change

As indicated above, an alternative potential source of forecast error are structural changes that alter the coefficients in the implicit model used by the OECD to forecast inflation. In particular, even if the OECD correctly forecast the exogenous variables, forecast errors might arise if there are changes in the responsiveness of inflation to these factors.

To test for this possibility, we regressed the forecast errors on the observed values of import prices, productivity growth, unemployment and the output gap. As suggested by the decomposition shown above, the parameters in these regressions are intended to represent estimates of the extent to which the actual parameters prevailing over the 1990s differ from those implicitly used by the OECD in making their inflation forecasts. As before, due to the paucity of observations, we used a set of bivariate regressions to assess the possibility of structural change rather than a single multivariate regression. Thus, these results, which are presented in Table 6, are again meant primarily to be illustrative.

Similar to the absence of a "smoking gun" in the correlations between forecast errors of inflation and forecast errors of the explanatory variable, there is no dominant parameter change evident in these results. However, in most countries, there are some coefficients that are suggestive of structural changes in the inflation process. In Australia, for example, the significant intercept coefficients provide some evidence of a shift in the average rate of both price and wage inflation over the 1990s, which could be interpreted as being associated with a decline in the NAIRU. The intercept term is also statistically significant for the price inflation errors in the United States, but given the absence of a significant effect on average wages and the evidence pointing to an increase in the response of wage growth to productivity growth, these results are more suggestive of a structural change involving the mark-up rather than of a fall in the NAIRU.

For other countries, the results are more difficult to interpret. For the United Kingdom, for example, there is no evidence of structural change in the parameters influencing price inflation, but substantial evidence of structural change in how wages are determined, although no single parameter dominates the results. In Japan, prices appear to have become more sensitive to the output gap and to import prices, while for wages, the intercept term points to a downward shift in the underlying pace of wage growth. For Spain, the estimates suggest that wage growth has become more responsive to productivity growth, which in turn has had a larger effect on price inflation. Finally, the results for Sweden point to a greater cyclical response in wages and a greater sensitivity of price inflation to import price shocks, while for Switzerland, there is evidence of a lower response of price inflation to changes in productivity as well as an increased sensitivity to changes in unemployment.

4. Structural changes, supply shocks and forecast errors

a. Some salient features of developments in the 1990s

As a preliminary to developing our own model-based tests of structural change, we first attempted to get a "feel" for such changes by looking at actual developments in indicators of activity and inflation (Table 7). In particular, we looked for evidence that the observed changes in price or wage inflation did not correspond to predictions based on standard Phillips curve relationships and for instances of changes in relative demand pressures in product and labour markets.

As indicators of resource utilisation, we use the log difference between actual and potential GDP (Gap)[16] as calculated by the OECD and the deviation of actual unemployment from our own estimates of the NAIRU (U*-U),[17] both measured as five-year averages for the three subperiods shown in the table. As indicators of inflation, we used the private consumption deflator and compensation per employee in the business sector; for both indicators, Table 7 includes two measures: the average rate of inflation for each five-year period (Δpc and Δw) and the change in the rate of inflation during each subperiod ($\Delta(\Delta pc)$ and $\Delta(\Delta w)$). The last two columns of the table give a sense of how the actual movements in the data correspond to the simple bivariate correlations implicit in simple Phillips curve relationships. In particular, a positive (negative) output gap might be expected to be associated with a rising (falling) rate of price inflation. Thus, a positive sign in the penultimate column ($\Delta(\Delta pc)$/Gap) of the table would indicate that the change in inflation is consistent with the hypothesised relationship, while a negative sign signals either that inflation has increased in a period when the output gap was negative or that inflation has fallen despite a positive gap. To identify possible asymmetries, the ratios are written in italics for periods of negative output gaps. The last column of the table contains a corresponding measure of the relationship between changes in the rate of wage inflation and labour market slack ($\Delta(\Delta w)$/(U* - U)). As with the product market measure, the ratio is constructed so that a negative sign is indicative of an unusual development, such as wage inflation falling (rising) when actual unemployment is below (above) the estimated NAIRUs.

For all countries, the early 1990s is the principal period of disinflation. Although declining rates of inflation are not surprising given that six of the eight countries recorded negative output gaps during this period, the degree of disinflation of both wage and price inflation is unusually large. This is particularly noticeable for the United Kingdom, Australia and Switzerland, but the decline in wage inflation in Spain and Sweden is also worth noting. Apart from the size of disinflation in these countries, the only "surprises" during the early 1990s were that both wage and price inflation declined in Japan and Spain, despite excess demand in both product and labour markets. For Japan, however, the OECD's estimate of the output gap is subject to a high degree of uncertainty, while in the case of Spain, the high rate of measured unemployment makes it difficult to derive a reliable measure of labour market slack, particularly for the 1990s, when measures deregulating the labour market were introduced.

While the early 1990s stand out as a period of disinflation with relatively few surprises, the second half of the 1990s contain several surprises but a relatively low degree of disinflation.[18] In terms of country-specific developments, the decline in price inflation in a period of excess demand in the product market is a main area of surprise for the United States, whereas wage developments are largely in line with predictions of a Phillips curve style relationship.[19] For Japan, the movements during 1995-99

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16 It would have been preferable to use rates of capacity utilisation as a complementary or alternative indicator, but comparable utilisation measures are only available for some of the countries.

17 Estimates of a time-varying U* were derived by applying a Kalman filter to an Okun's law relationship using the OECD's estimates of potential GDP. Thus for U<U* or excess demand in the labour market, the deviation will be positive and comparable to the Gap measure of conditions in the output market.

18 For several countries (notably Japan and Switzerland), the low degree of disinflation is a "non-surprise" given that the rate of price inflation had already fallen to very low levels by the end of the 1990-94 period. In other countries (the United Kingdom, Canada and Australia), the decline in wage inflation may have slowed as actual rates of unemployment came closer to, or fell below, the estimated NAIRUs.

19 Gordon (1998) makes a similar point.

largely correspond to the predictions of the Phillips curve: price inflation declined in the presence of a large and widening output gap while wages decelerated in response to rising unemployment. Nonetheless, the fact that the rate of disinflation did not increase as the output gap widened is surprising and might indicate a change in firms' price setting behaviour.

In the case of the United Kingdom, the principal period of disinflation was obviously the first half of the 1990s: price inflation declined by 5.5 percentage points despite a relatively small output gap, while wage inflation declined by almost 7 percentage points, even though the actual rate of unemployment was only slightly above the estimated NAIRU. Developments during the second period are similar to those observed for the United States. The main surprise is that price inflation declined further despite a positive output gap. In contrast, the acceleration of wages is in line with unemployment falling below the NAIRU.

For Canada, the changes observed in inflation and activity measures are in line with predictions from the Phillips curve for both the first and the second half of the 1990s, though the smaller degree of deceleration during the second half is worth noting.[20] As in the United Kingdom, the early 1990s were the principal period of disinflation in Australia. However, even though there were signs of excess demand in both product and labour markets during 1995-99, prices and wages continued to decelerate. The fact that wages also continued to decelerate in Spain, Sweden and Switzerland is less surprising, as actual rates of unemployment exceeded the NAIRU. Similarly, the continued decline in price inflation is in line with predictions from the Phillips curve as all three countries recorded relatively large output gaps.

b. Price and wage model forecast errors

We next turn to an evaluation of forecast errors from our own price and wage equations. Our procedure is as follows:[21]

- First, we estimated wage and price equations for the period 1960-90 using an error correction model allowing for wage-price feedbacks in both levels and first differences;[22]
- Second, when the error correction term was insignificant, we reestimated the equations and tested the homogeneity constraint implied by the assumption that wage and price inflation are I(1) processes;
- Third, when the price-wage feedback term in first differences was also insignificant, we estimated reduced-form price equations with and without the homogeneity constraint;
- Finally, the most satisfactory versions were used to make forecasts for the period 1991-99 and reestimated over the full sample period (1960-99) in an attempt to identify parameter changes as possible sources of forecast errors.

Price equations

The price equations estimated under the first step mentioned above were specified as:[23]

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[20] The behaviour of price inflation in Canada in part depends on the measure used. While the change in the private consumption deflator was largely constant during 1995-99, inflation measured by the GDP deflator declined by 0.6 percentage points. In contrast, during 1990-94 the GDP deflator decelerated by only 2 percentage points, compared with more than 3 percentage points for the consumption deflator.

[21] Before adopting this procedure, we tested the dynamic properties of the relevant variables, using an augmented Dickey-Fuller test. With the exception of Switzerland, the tests suggested that price inflation is an I(1) process, while wage inflation is found to be I(1) in all countries save Australia. As might be expected, the output gap is I(0) while actual unemployment tends to be I(1). However, when measured as a deviation from the NAIRU, unemployment also becomes I(0).

[22] For some countries, we had to derive our own estimates for the output gap in the 1960s, using a quadratic trend and the rate of unemployment, and then combine them with the output gaps published by the OECD.

[23] It is debatable whether the GDP deflator or the consumption deflator should be used in measuring real compensation. Since the price equation refers to consumer price inflation, we used the consumption deflator. We also estimated an alternative set of regressions, using changes in unit labour costs instead of compensation per employee. Significant coefficients of the expected positive sign were only found for Canada and Sweden, but only when the sample period was extended to include

$$\Delta pc = \alpha + \beta Gap_{-1} + \gamma\Delta pc_{-1} + \eta\Delta pc_{-2} + \varsigma\Delta w_{-1} + \lambda((w - pc) - q)_{-1} + \phi(\Delta pm - \Delta pc_{-1}) + \varepsilon \quad (9)$$

Where *pc* denotes the (log) private consumption deflator, *w* (log) compensation per employee, *q* (log) output per employed person, Gap the output gap, and *pm* (log) import prices. Subscripts refer to lags in years and Δ is the first-difference operator.

In equation (9), the coefficient on the error correction term (λ) is expected to be positive if a decline in real wages relative to productivity puts downward pressure on prices.[24] However, when λ was negative or insignificant, the price equation was specified as:

$$\Delta pc = \alpha + \beta Gap_{-1} + \gamma\Delta pc_{-1} + \eta\Delta pc_{-2} + \varsigma\Delta w_{-1} + \phi(\Delta pm - \Delta pc_{-1}) + \varepsilon \quad (10)$$

As the unit root tests frequently indicated that price inflation was I(1), we also estimated a variant of (10) constraining the coefficients on lagged rates of price and wage inflation to sum to unity (ie with $\gamma+\eta+\zeta=1$ imposed). Finally, in those cases where we were unable to identify any significant wage-price feedbacks, we dropped the lagged wage term for both the unconstrained and constrained versions of the model.

Before discussing the results from our preferred specifications, it is worth pointing out that estimates of (9) provided little support for the hypothesis that firms increase prices in response to a rise in real unit labour costs (or a fall in the mark-up). For five countries we obtained the "wrong" sign and in two cases the coefficient on real unit labour costs was insignificant, albeit positive. Only Sweden exhibited a positive and significant coefficient, although this result was limited to the longer sample period.

The estimates obtained from equation (10) also provided little evidence of wage-price feedbacks. In particular, statistically significant coefficients on the lagged wage terms were evident only for Japan, Switzerland and Spain.[25] In the case of Japan, the addition of lagged wage changes was a significant improvement compared with the results obtained from the price equations excluding this feedback. The diagnostic statistics were better and, above all, the tendency to overpredict was much lower. For Switzerland and Spain, we also obtained significant coefficients for the lagged change in wages, and the forecasts derived from this specification were fairly accurate. Overall, however, the results were less satisfactory than those based on the reduced-form price equations. Particularly for Spain, this finding is somewhat surprising as the deceleration of wages has been more pronounced than in most other countries; yet, lagged wage changes seem to have had only a marginal impact on the path of consumer prices.

As a result, with the exception of Japan, Table 8a only presents estimates for the reduced-form price equations. However, because there are major differences in their ability to forecast, we report results for both the unconstrained and the constrained versions of the model.

Turning first to the estimates for the unconstrained equations, the coefficients on the output gap range from 0.2 to 0.6 and are statistically significant for six of the eight countries. Except for Japan, where lagged changes in wages dominate lagged price changes, all the coefficients on the one-year lagged inflation term have significant t-values. In contrast, most of the coefficients on the two-year lagged inflation term are insignificant, though none of the t-values is less than unity. Changes in relative import prices, which were included as a proxy for supply shocks, are significant for all countries, and in several cases the coefficients are close to the import share of GDP. The autocorrelation statistics are not indicative of specification errors and the R^2s range from 0.7 to 0.95.

For the constrained version, the coefficient on the output gap is higher for five of the countries, but is somewhat lower for Switzerland and, especially, Australia. Most of the coefficients on the lagged inflation terms are negative and imply a slightly greater influence of one-year lagged inflation as compared with the unconstrained version; in the case of Japan, the influence of lagged changes in

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the 1990s. Finally, we experimented with longer lags on the determining variables in (9) but, with the exceptions of Australia (two-year lag on the output gap) and the United Kingdom (both current and one-year lagged changes in import prices significant), the lag specification shown in equation (9) worked best.

24 More particularly, since wages decelerated faster than consumer prices during the 1990s, leaving out the feedback from wages to prices might be a source of overpredicting price inflation. Moreover, as found by Brayton et al (1999) and Gali et al (2000), firms' pricing decisions seem to have been affected by deviations of real unit labour costs from their long-run trend.

25 Replacing compensation per employee with a measure of unit labour costs did not change this result.

real wages has the right sign but is not significant. The influence of relative import prices is relatively stable, the largest change being observed for Switzerland.

According to the F-statistics shown in the penultimate row of the table, the homogeneity constraint is rejected for five of the countries, most convincingly for those (Switzerland, the United States and Australia) where Δpc is (or is close to being) an I(0) variable. Nonetheless, except for the United Kingdom, the constrained version produces smaller forecast errors than the unconstrained version despite the fact that the constrained version assigns larger weights to past rates of inflation and that inflation decelerated during the 1990s.[26] This finding stands in contrast to the results reported by Stock and Watson (1999) who find that the unconstrained version of their model generates better forecasts for US inflation in the more recent years when inflation was low and close to being an I(0) variable.

Looking more specifically at the size of the forecast errors, Australia, Spain, and Canada show the largest cumulative errors for simulations based on the unconstrained equation, while those for Switzerland, the United States and the United Kingdom are quite small.[27] Together with Japan, these are also the only countries for which inflation is actually underpredicted for the 1990s once the homogeneity constraint is imposed.[28]

The parameters obtained when extending the sample period to 1999 are presented in Table 8b. For neither the unconstrained nor the constrained version of the price equation can the hypothesis of parameter stability be rejected for any country when the test is applied to all the parameters jointly.[29] Nonetheless, there are notable changes in the point estimates of the coefficients which may help to explain the forecast errors discussed above:

- For most countries, the intercept term declines when the sample is extended to include the 1990s. The decline is most pronounced for Australia and Spain; ie the two countries with the largest prediction errors in Table 8a;
- For both Japan and Australia, the influence of product market conditions declines significantly when the 1990s are included. In the case of Japan, this implies that the rise in the output gap since 1997 has had a smaller disinflationary (or deflationary) impact than suggested by historical patterns. For Australia, the rise in actual relative to potential GDP during the second half of the 1990s had a smaller impact on the rate of inflation than the historical estimates would have implied. Spain also experienced a decline, albeit small, in the gap coefficient while Canada shows a slight increase. For the other four countries, there were virtually no changes in the parameter on the output gap;
- As discussed earlier, there is some ambiguity with respect to the influence of lagged inflation in a low-inflation regime. The results obtained by extending the sample period to include the 1990s "straddle the fence" between the opposing views, as the sum of the coefficients on the lagged inflation terms change only marginally. The main exceptions are Australia and Switzerland, for which inflation persistence increases somewhat;

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[26] A marked drop in the intercept terms in the constrained equations accounts for a substantial part (largest for Switzerland and Australia and lowest for Canada and Spain) of the smaller overprediction for the 1990s. However, since we know little about the causes of this change, its contribution should be regarded as a quantification of our surprise rather than as an explanation.

[27] We also calculated forecast errors based on dynamic simulations. While this method increased the size of the errors, the ranking of the countries did not change.

[28] Using the unconstrained equation, it appears that the forecast errors for the United States and Australia were evenly distributed between the two halves of the 1990s, while for Japan, the United Kingdom, Canada and Spain the overpredictions were most pronounced during the first half, when the speed of disinflation was also most pronounced. The equation for Switzerland underpredicted inflation during the first half but overpredicted during the second half, while the errors for Sweden are dominated by an especially large miss in 1991. We also tested whether the tendency to overpredict increased as the rate of inflation declined by regressing the forecast errors on the actual rates of inflation. The hypothesis is confirmed for the whole sample period. However, when the regression is confined to the out-of-sample prediction errors, we only obtain significant and negative coefficients for Sweden and Switzerland.

[29] On the assumption that structural changes have taken place but that the source of the changes is not known, we also estimated the price equations allowing the intercept term to differ after 1990. For the unconstrained model, the coefficients on the dummy variable were always negative but statistically significant only for Australia, though the coefficients for Japan and Canada also obtained relatively high t-values. The coefficients for the constrained model were also mostly negative but never came close to being significant.

- The hypothesis that the parameters of the price equation may be regime-dependent is also not supported by the coefficients obtained for changes in relative import prices in the extended sample. Except for Sweden, for which the pass-through actually increases,[30] the coefficients remain remarkably constant, suggesting that the influence of changes in relative import prices in the 1990s was little different from that of earlier decades.

To sum up, when extending the sample period to 1999, the cumulative within-sample forecast errors for both the unconstrained and the constrained equations decline to only marginal levels and, except for Sweden, the principal parameter change contributing to this improvement is the decline in the intercept term. For the two countries with the largest forecast errors when using the shorter sample (Australia and Spain), the lower intercept terms would imply a cumulative reduction in the inflation forecasts of 4-8 percentage points. However, while these results are indicative of structural change, there is little information as to what might have caused this decline.

Wage equations

Similar to the procedure used for estimating price equations, we initially attempted to explain wage changes using a specification with an error correction term:

$$\Delta w = \alpha + \beta(U - U^*)_{-i} + \delta\Delta(U - U^*)_{-i} + \lambda((w - pc) - q)_{-1} + \sum_i \phi_i \Delta w_{-1} + \sum_j \varphi_j \Delta pc_{-j} + \varepsilon \qquad (11)$$

with w denoting log compensation per employee, U-U* the deviation between actual unemployment and the NAIRU, pc the log private consumption deflator, q log productivity, Δ the first-difference operator, i = 1, 2 and j = 0, 1. On the assumption that employees will push for higher wages when their income share declines, the coefficient on the error correction term (λ) is expected to be negative. However, when λ was either positive or insignificant, (11) was reestimated without the error correction term:

$$\Delta w = \alpha + \beta(U - U^*)_{-1} + \delta\Delta(U - U^*)_{-1} + \sum_i \phi_i \Delta w_{-i} + \sum_j \varphi_j \Delta pc_{-j} + \varepsilon \qquad (12)$$

Moreover, since the gap between actual unemployment and the NAIRUs is an I(0) variable and wage inflation in most countries is I(1), we also estimated an equation with consistent dynamic structures by forcing the coefficients on the lagged wage and price changes to sum to unity (ie with $\Sigma\phi_i+\Sigma\varphi_j=1$).[31] As with the constrained price equation, the dynamics of the inflation process are reflected in the lagged dependent variable while the lagged change in real wages captures potential feedbacks between price and wage inflation. However, in contrast to the price equations, for which we were able to use fairly similar specifications for all countries, it turned out to be far more difficult to find comparable wage equations. In several cases it was necessary to extend the lag structure on price changes in order to obtain specifications that did not suffer from autocorrelation. Moreover, for three countries, actual unemployment rather than the unemployment gap seemed to be the best indicator of labour market conditions.

The results obtained from applying the same four-step procedure as for price inflation are displayed in Tables 9a (shorter sample period) and 9b (extended sample period). In general, the cumulative forecast errors tend to be of the same size as for prices, though three differences compared with the price equations are worth noting. First, for about half the countries, the constrained equation does not generate lower forecast errors, even though the constraint is never rejected. Second, wage inflation for the 1990s is underpredicted in half the cases, while for price inflation, positive forecast errors were recorded for nearly all countries. Third, when the sample period is extended to include the 1990s,

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30 This seems to be the principal reason for the lower forecast error for Sweden. However, the autocorrelation coefficients suggest that the equation is misspecified when the data sample includes the 1990s.

31 On the assumption that wage demands may be influenced by expected developments in labour productivity, we also estimated equations including changes in output per person employed, smoothed by an HP filter. In all cases, the productivity term entered with a positive coefficient and for about half the countries it was also significant. However, in most cases, the overall fit was inferior to those of (11) and (12) and coefficients that were significant in (11) and (12) were less precisely estimated. The only exceptions were Japan and Switzerland, for which the addition of changes in productivity substantially improved the fit. However, as discussed below, even better results for these two countries were obtained when the influence of productivity developments was captured through the error correction term.

parameter stability is rejected (or is close to being rejected) for about half the countries; in contrast, the coefficients of price inflation were mostly stable.

Turning to the results for individual countries, the wage equation for the *United States* slightly underpredicts wage inflation in the 1990s. The homogeneity constraint is not rejected and the coefficients not only confirm the long lag structure found in other analyses of US wage inflation but also suggest that current wage inflation tends to rise in response to past declines in real wages (see also Gordon (1998)). When the equations are extended to include the 1990s, there are signs of parameter instability. In particular, the lag structure appears to have become shorter while the response of current wages to past changes in real wages increases somewhat. The intercept term also increases, perhaps indicative of stronger productivity growth and associated real wage gains, while the sensitivity to the unemployment gap is unchanged.

The unconstrained wage equation for the *United Kingdom* has the highest prediction errors in our sample, though the errors are concentrated in the first half of the 1990s, when the deceleration of wages was most pronounced. The imposition of the homogeneity constraint leads to a marked decline in the intercept term and a stronger reaction to changes in labour market conditions, significantly improving the accuracy of the predictions. Unlike the equations for the United States, the coefficients appear to be stable when the sample period is extended to 1999, although the tendency to overpredict remains. Another feature of the UK equation, and a possible source of the overpredictions, might be that the coefficient on the one-year lagged inflation rate exceeds unity. In a period of rapidly decelerating prices, this tends to keep wages growing faster than they otherwise would have done.

For both *Canada* and *Australia*, the actual rate of unemployment (lagged one year) produced better results than the unemployment gap. Nevertheless, a specification with U rather than U-U* raises several questions. First, such a specification implicitly assumes that U* has been constant over the period, in contrast to the pattern suggested by the NAIRUs calculated from the OECD output gaps. Second, there are issues about the correct modelling of the dynamic structures. If wage inflation is an I(1) variable, the unconstrained version of equation (12) would be the correct specification, assuming that Δw and U are cointegrated. Conversely, if wage inflation is I(0), an error correction term should be included (as in equation (11)) to ensure dynamic consistency.

For *Canada*, the empirical evidence favours the first hypothesis. Even though wage inflation is close to being I(0), the error correction term was insignificant and had the wrong sign.[32] In addition, equation (12) yields smaller cumulative forecast errors than for any other country. Moreover, there is no sign of parameter instability when the sample is extended to include the 1990s, possibly suggesting that the equation is capturing a cointegrating relationship between wage inflation and the rate of unemployment that is suppressed in the constrained version.

The estimates for *Australia* are considerably more problematic.[33] First, the dynamic response of wages to prices is highly erratic, with a large positive coefficient on the first lag of inflation and a large negative coefficient on the second lag. This pattern produces large prediction errors for individual years, with both the constrained and unconstrained versions of the model underpredicting wage inflation in the 1990s by a cumulative 19 percentage points.[34] Second, including an error correction term, as in (11), helped only little. The volatile dynamic structure remained and the coefficient on the error correction term was insignificant, though of the right sign. Third, while parameter stability cannot be rejected when the sample is extended to include the 1990s, the coefficient on the unemployment rate *does* decline substantially, and this change alone reduces the predicted wage changes in the 1990s by a cumulative 15½ percentage points. However, while the parameter shift is significant, it raises the question as to why, in a period of deregulation and a more flexible labour market, wage earners in Australia should have become less sensitive to changing market conditions. Indeed, the

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32 When the error correction terms were included together with (U-U*), we obtained a highly significant coefficient of the correct (negative) sign. However, U-U* still provided no explanatory power.

33 We attempted to compare our estimates with those presented in Gruen et al (1999). However, their results are based on four-quarter changes of consumer prices and unit labour costs and the time profile of their estimates of the NAIRU is quite different from ours, making such comparisons difficult.

34 We experimented with various lag structures for both wage and price inflation, but those shown in the tables produced the best diagnostic statistics as well as the smallest overall prediction errors.

downgrading of the Arbitration system and the greater role of bargaining at the firm level would seem to point to a *greater* sensitivity of wages to labour market developments.

Sweden is another country for which wage inflation in the 1990s is underpredicted and, again, the main source of the prediction errors appears to be parameter instability, with the largest shifts observed for the intercept term and the coefficients on the level of and changes in the unemployment gap. For the constrained version, these parameter changes alone “explain” 75% of the cumulative underpredictions. There is also some decline in the influence of current price inflation and lagged wages, but this changes the predictions only marginally.

The wage equation for Spain contains a relatively large autoregressive component which, combined with the sharp deceleration of wages in the 1990s, appears to be the main source of the overpredictions. The feedback from prices to wages is relatively moderate and, except for the intercept term, the parameters remain rather stable for the longer sample period. Although the homogeneity constraint cannot be rejected, the constrained version of (12) predicts less well and produces less satisfactory diagnostic statistics.

We next turn to the two countries for which the specification including an error correction term (equation (11)) produced the best results. In the case of Switzerland, adding the error correction term leads to a substantial improvement compared with the properties of equation (12), including a marked decline in autocorrelation and a dynamic structure that looks more plausible. Moreover, the unemployment gap is highly significant, the R^2 rises to 0.87, the cumulative prediction error is only 0.7 percentage points, and the parameter estimates change little when the sample is extended to include the 1990s.

At first glance, the wage equation for Japan also looks plausible when allowing for price-wage feedbacks. The error correction term is highly significant and the diagnostic statistics are satisfactory. Moreover, because real unit labour costs increased by almost 5% during the first half of the 1990s, significantly depressing wage demands later in the decade, wage inflation is underpredicted by a cumulative 7½ percentage points, compared with an overprediction of 14½ percentage points when using equation (12).[35] Yet, the addition of real unit labour costs is not entirely satisfactory. Even though the hypothesis of parameter stability cannot be rejected when the sample period is extended to include the 1990s and the within-sample forecast error is reduced to less than 3 percentage points, there are clear signs of parameter instability, as both the intercept term and the coefficient on unemployment become insignificant. Moreover, the coefficient on the error correction term increases (in absolute terms) to 24.7, almost twice the value obtained for the shorter sample.

5. Conclusions

An important and welcome feature of economic developments in the 1990s was the extent to which inflation fell in an environment of high or rising economic activity and tightening labour markets. At the same time, the fact that most forecasters underestimated the degree of disinflationary pressure during the decade raises questions about the appropriate stance of monetary policy. If the forecast errors mainly reflect the effects of favourable supply shocks but policymakers attribute them to permanent structural changes, there is a risk that monetary policy will be too expansionary and that the disinflationary gains will be lost. Conversely, if the unexpected declines in inflation are due to irrevocable changes in the inflation process but policymakers interpret them as only transitory, policies might be kept too tight, with a risk that potential gains in output and employment will not be realised.

In this paper, we first document the extent to which inflation was overpredicted in the 1990s, using a sample of eight industrialised countries and forecasts produced by the OECD as well as by our own models. According to both sources, the overprediction of inflationary pressures has been most pronounced for prices, whereas wage gains have been more in line with predictions based on traditional Phillips curves. This is an important finding as it suggests that the source of the forecast

[35] The change in the sign of the prediction errors might be related to the fact that when including real unit labour costs as a determinant, the measured rate of unemployment (rather than its deviation from the NAIRU) appears to be the appropriate measure of labour market slack.

errors is mainly to be found in firms' price setting behaviour. At the same time, it leaves open the question of whether the implied reduction in firms' mark-ups reflect favourable supply shocks, such as lower relative import prices and one-time productivity gains, or a more permanent erosion of firms' pricing power due to other factors, such as globalisation and deregulation of product markets.

Consequently, the paper attempts to identify the sources of the apparent change in firms' price setting behaviour, relying again on forecasts made by the OECD and our own price and wage equations, and using a wide range of econometric and non-econometric tests. While we do find some evidence of structural shifts in the inflation process, in only a few cases are we able to identify the source of these shifts. Moreover, the significant changes that we find differ widely across the eight countries in their nature as well as their size. For instance, for the United States and Australia it appears that higher productivity gains are a main source of the tendency to overpredict price inflation. In contrast, the overpredictions of inflation in Japan and Switzerland seem to reflect movements in real unit labour costs (or mark-ups), which traditional Phillips curves tend to ignore. In the case of Sweden, favourable import price shocks appear to be the main source of overpredicting consumer price inflation while for the three remaining countries (the United Kingdom, Canada and Spain) we were unable to find any dominant cause.

In some ways, it is not surprising that we are unable to find a unique explanation of these forecast errors. While disinflation and monetary policies aimed at price stability were common to all countries during the 1990s, the accompanying cyclical developments progressively diverged. In the four English-speaking countries, economic activity tended to exceed earlier estimates of potential output levels, while continental Europe experienced a slow and moderate recovery from the recession in the early 1990s and Japan moved deeper into recession. The processes of deregulation also differed. In the United States, few regulations were left by the start of the 1990s, and the other English-speaking countries either took major steps to deregulate labour and product markets during the 1990s or had already done so during the late 1980s. In continental Europe, by contrast, progress has been much slower, and even though many rigidities have been removed, these changes have probably not yet affected the inflation process in a significant way. Finally, in Japan, major reforms have been announced but the implementation of reform measures has barely started. As a result, the influence of globalisation and developments associated with the new economy on firms' price setting behaviour is likely to have differed widely, being most pronounced in those countries that have gone furthest in liberalising their markets.

In addition to our inability to clearly identify and explain the sources of the changes we find, the paper leaves a number of other questions unresolved, some of which are also relevant to policies. One puzzling result is that the homogeneity constraints on the price equations are often rejected, and yet the constrained versions produce more accurate forecasts. Related to this, and of potential importance to forecasting procedures, we were unable to find any evidence of the regime-dependent changes in the parameters of the inflation process highlighted in recent empirical work. In particular, neither the OECD forecasts nor our own estimates suggest that economic agents tend to ignore inflation when it is very low or that the pass-through of import prices and other costs is smaller in a low-inflation environment. Another puzzle is that even though our estimated price equations tend to overpredict inflation in the 1990s, the parameters appear to be stable. In contrast, several of our wage equations show signs of parameter instability but produce forecasts that are less biased. Whatever the source of this inconsistency and despite the evidence pointing to changes in firms' price setting behaviour as the likely source of the forecast errors, it implies that in order to better understand the inflation process in the 1990s, it is important to look at both the price and the wage components of the Phillips curve framework.

Against this background, it is also not surprising that the empirical evidence we present does not clearly favour any of the alternative theoretical explanations of recent inflation performance discussed in Section 2, either for the eight countries as a group or for any individual country. In particular, even though parameter shifts and specific sources of forecast errors have been identified in a number of cases, we are unable to distinguish between one-time shock effects and permanent changes in the inflation process. This is unfortunate and disappointing from a policy point of view; however, it should not be regarded as a major surprise, considering that most of our evidence is based on only nine years of observations while the debate about shocks versus permanent changes is at least 25 years old.

Table 1

Forecast errors for US inflation

Year	OECD (PCE deflator)			FOMC (CPI)		
	Forecast	Outcome	Difference	Forecast[1]	Outcome	Difference
1991	4.6	3.8	0.8	4.1	3.0	1.1
1992	3.9	3.1	0.8	3.5	3.1	0.4
1993	3.2	2.4	0.8	3.0	2.7	0.3
1994	2.8	2.0	0.8	2.7	2.6	0.1
1995	3.1	2.3	0.8	3.1	2.7	0.4
1996	3.4	2.1	1.3	3.1	3.1	0.0
1997	2.3	2.0	0.3	2.9	1.9	1.0
1998	2.4	0.9	1.5	2.7	1.5	1.2
1999	1.7	1.6	0.1	2.2	2.6	– 0.4

[1] Based on the midpoint of the central tendency.

Table 2

Forecast errors for price inflation in the 1990s

Selected countries, 1991-99, annual data

	United States	Japan	United Kingdom	Canada	Australia	Spain	Sweden	Switzerland
Summary statistics								
MAE	0.80	0.53	0.58	0.67	1.36	0.57	1.23	1.00
RMSE	0.89	0.63	1.03	0.85	1.45	0.70	1.42	1.19
Bias measures								
Mean	0.80[2]	0.26	– 0.12	0.58[1]	1.36[2]	– 0.32	0.33	0.32
U^M	0.81	0.17	0.01	0.46	0.88	0.22	0.06	0.07
% > 0	1.00	0.78	0.33	0.89	1.00	0.33	0.78	0.78
Efficiency								
α	0.29	0.25	1.69	0.19	0.50	– 0.04	1.06	2.54[2]
$1 - \beta$	0.17	0.00	– 0.53	0.17	0.26[1]	– 0.07	– 0.20	– 0.95[2]
Joint F-test	17.30[2]	0.70	1.19	3.69[1]	46.16[2]	1.08	0.60	5.64[2]
Serial correlation								
Q(1)	6.55[2]	0.01	0.17	0.46	2.54	0.43	1.70	2.00
Q(2)	9.46[2]	3.78	0.36	0.65	2.68	0.65	3.84	2.19

Definitions*:* Price inflation is measured as the percentage change in the personal consumption deflator. Forecast errors are measured as forecast less outcome; ie a positive (negative) error indicates that inflation has been over-(under-)predicted. Notation: MAE = mean average forecast error; RMSE = root mean squared errors; mean = average forecast error; U^M = the proportion of the mean squared error due to the mean error; % > 0 = the proportion of years in which the OECD overpredicted inflation; α and $1 - \beta$ = coefficients obtained when regressing forecast errors for year t on an intercept term and the predicted inflation rate for year t; joint F-test tests the hypothesis $\alpha = 0$ and $\beta = 1$; Q(1) and Q(2) = Ljung-Box Q-statistics for 1 and 2 lags respectively.

[1] and [2] denote 90 and 95% levels of significance.

Table 3

Forecast errors for wage inflation in the 1990s

Selected countries, 1991-99, annual data

	United States	Japan	United Kingdom	Canada	Australia	Spain	Sweden	Switzerland
Summary statistics								
MAE	1.13	1.65	1.56	1.14	1.47	2.16	1.43	0.85
RMSE	1.24	1.87	1.85	1.56	1.71	2.49	1.64	1.08
Bias measures								
Mean	0.48	1.65[1]	0.39	0.34	1.21[1]	– 0.04	0.19	0.32
U^M	0.15	0.78	0.04	0.05	0.50	0.00	0.01	0.09
% > 0	0.55	1.00	0.56	0.89	0.89	0.67	0.44	0.56
Efficiency								
α	– 0.94	1.36	– 1.28	– 1.14	– 2.42	6.61[1]	– 2.69	0.69
$1 - \beta$	0.35	0.11	0.32	0.45	0.82	– 1.36[1]	0.64	– 0.12
Joint F-test	0.78	12.73[1]	0.41	0.45	7.15[1]	10.85[1]	1.19	0.51
Serial correlation								
Q(1)	0.00	1.18	3.10	0.24	0.57	2.34	0.24	0.19
Q(2)	0.00	2.59	3.12	0.32	0.97	2.51	0.13	0.28

Definitions: Wage inflation is measured as the percentage change in compensation per employee. Forecast errors are measured as forecast less outcome; ie a positive (negative) error indicates that inflation has been over-(under-)predicted. Notation: MAE = mean average forecast error; RMSE = root mean squared errors; mean = average forecast error; U^M = the proportion of the mean squared error due to the mean error; % > 0 = the proportion of years in which the OECD overpredicted inflation; α and $1 - \beta$ = coefficients obtained when regressing forecast errors for year t on an intercept term and the predicted inflation rate for year t; joint F-test tests the hypothesis α=0 and β=1; Q(1) and Q(2) = Ljung-Box Q-statistics for 1 and 2 lags respectively.

[1] Denotes a 95% level of significance.

Table 4

Exogenous variables: means (μ) and average forecast errors (ε)

1991-99, based on data as of June 2000

Country	Import prices (%Δ)		Productivity (%Δ)		Unemployment (per cent)		Output gap (per cent)	
	μ	ε	μ	ε	μ	ε	μ	ε
United States	– 0.9	1.6	1.8	– 0.6[2]	5.8	– 0.1	– 0.2	0.8
Japan	– 2.6	3.9	0.9	1.1[2]	3.1	– 0.3[3]	– 0.6	0.0
United Kingdom	0.3	2.0	1.9	0.3	8.2	– 0.1	– 0.6	– 1.2[3]
Canada	1.7	0.0	1.2	0.4	9.7	– 0.3	– 2.3	0.9
Australia	0.5	0.7	2.4	– 1.0[3]	9.0	– 0.2	– 0.4	– 0.7
Spain	2.5	0.1	1.5	0.1	20.0	– 0.1	– 0.5	0.4
Sweden	2.3	0.3	2.5	– 1.1[3]	6.7	– 0.7[2]	– 2.4	1.4[2]
Switzerland	– 0.8	3.0[3]	0.4	0.9[2]	3.7	– 0.5	– 2.3[1]	– 0.8[1]
Average	*0.4*	*1.5*	*1.6*	*0.0*	*8.3*	*– 0.3*	*– 1.2*	*0.1*

[1] Output gap forecasts for Switzerland are only available from 1996 to 1999. [2] and [3] denote 90% and 95% levels of significance.

Table 5

Bivariate regressions of forecast errors[1]

1991-99, based on data as of June 2000

Country	Import prices		Productivity		Unemployment		Output gap		Δpc or Δw	
	Δpc	Δw	Δpc	Δw	Δpc	Δw	Δpc	Δw	Δpc	Δw
United States	0.20[4]	0.05	– 0.50[3]	0.44	– 0.08	0.44	– 0.08	0.60	0.31	0.61
Japan	0.05[4]	0.09	– 0.68[3]	0.57[3]	0.12	0.57[3]	0.12	0.12	0.15	1.34
United Kingdom	– 0.05	– 0.24[3]	– 0.76	– 0.23	0.87[3]	– 0.23	0.87[3]	– 0.23	0.11	0.35
Canada	– 0.07	– 0.16	0.31	0.31	– 0.19	0.31	– 0.19	0.32	0.32[3]	1.06[3]
Australia	0.12	– 0.04	– 0.63[3]	– 0.39	– 0.46	– 0.39	– 0.46	– 0.02	0.60[4]	0.84[4]
Spain	0.12	– 0.08	0.18	1.19	0.08	1.19	0.08	– 0.67	0.04	0.46
Sweden	0.18[4]	0.05	– 0.07	– 0.08	0.38	– 0.08	0.38	0.14	0.24	0.32
Switzerland	0.08	0.13	– 0.31	0.06	0.21	0.06	0.21	0.30[2]	0.50	0.41

[1] Coefficients are those obtained by regressing (with no intercept term) forecast errors for Δpc and Δw on forecast errors for the exogenous variables listed at the top, with the figures in the last column showing the coefficients from bivariate regressions between forecast errors for Δpc and Δw. [2] Based on only four observations. [3] and [4] denote 90% and 95% levels of significance.

Table 6

Bivariate regressions of forecast errors on outcomes[1]

1991-99, based on data as of June 2000

Country	Intercept[2]		Import prices		Productivity		Unemployment		Output gap		Δpc or Δw	
	Δpc	Δw	Δpc	Δw	Δpc	Δw	Δpc	Δw	Δpc	Δw	Δpc	Δw
United States	0.78[5]	0.37	– 0.06	0.22	0.05	– 1.23[5]	0.05	0.21	– 0.07	– 0.45	0.18	– 0.18
Japan	0.15	1.66[5]	– 0.06[4]	– 0.01	– 0.10	– 0.17	0.13	– 0.15	– 0.17[5]	0.01	0.26	0.56
United Kingdom	– 0.12	– 0.36	0.04	0.31[5]	0.69	1.39[4]	0. 10	1.07[5]	0.01	– 1.26[5]	– 0.03	0.70[4]
Canada	0.18	– 0.39	0.06	0.18	0.20	0.34	0.20	0.38	– 0.17	– 0.32	0.42[5]	0.23
Australia	1.32[5]	1.08[5]	– 0.07[4]	0.07	0.00	– 0.23	0.05	0.09	– 0.10	– 0.36	0.24	0.15
Spain	– 0.33	– 0.37	– 0.12	– 0.02	– 0.18	– 1.33[5]	– 0.04	0.12	– 0.01	– 0.68	– 0.20	– 1.17[5]
Sweden	1.19	1.19	– 0.21[5]	– 0.09	– 0.15	– 0.34	– 0.01	– 0.56[4]	0.36	0.42	0.15	0.14
Switzerland	1.85	– 3.34	– 0.15	– 0.15	0.66[5]	0.06	0.68[5]	0.46	0.35[3]	– 1.55[3]	– 0.26[5]	0.13

[1] Coefficients obtained by regressing forecast errors for Δpc and Δw on an intercept term and the exogenous variables listed at the top. [2] Coefficients are taken from a regression of the forecast errors on an intercept term and the output gap. [3] Regressions including the output gap are based on only four observations. [4] and [5] denote 90% and 95% levels of significance.

Table 7

Activity indicators and inflation in the 1990s

Country	Period	Gap	Δpc	$\Delta(\Delta pc)$	$U^* - U$	Δw	$\Delta(\Delta w)$	$\Delta(\Delta pc)/Gap$	$\Delta(\Delta w)/(U^* - U)$
United States	1985-89	0.7	3.6	0.9	0.3	4.1	– 0.8	1.3	– 2.7
	1990-94	– 0.7	3.2	– 2.6	– 0.4	3.9	– 2.6	*3.7*	*6.5*
	1995-99	0.6	1.8	– 0.7	0.3	3.5	2.5	– 1.2	8.3
Japan	1985-89	– 1.2	1.2	– 0.2	– 0.2	3.0	0.3	*0.2*	*– 1.5*
	1990-94	0.8	1.8	– 1.8	0.2	2.6	– 3.1	– 2.1	– 15.5
	1995-99	– 1.5	0.2	0.0	– 0.4	0.2	– 1.2	*0.0*	*3.0*
United Kingdom	1985-89	2.8	4.9	1.0	1.1	8.2	1.2	0.3	1.1
	1990-94	– 0.8	5.2	– 5.5	– 0.2	5.7	– 6.7	*6.9*	*33.5*
	1995-99	0.3	2.7	– 0.5	0.2	5.0	2.1	– 1.7	10.5
Canada	1985-89	2.0	4.0	0.5	0.9	5.5	– 0.6	0.2	– 0.7
	1990-94	– 2.7	2.7	– 3.2	– 1.3	3.2	– 4.5	*1.2*	*3.5*
	1995-99	– 1.3	1.3	0.0	– 0.7	3.1	– 0.2	*0.0*	*0.3*
Australia	1985-89	1.0	7.2	– 1.4	0.3	6.1	2.4	– 1.4	8.0
	1990-94	– 1.4	3.2	– 5.3	– 0.7	4.0	– 6.4	*3.8*	*9.1*
	1995-99	0.8	1.5	– 0.7	0.3	3.5	– 0.9	– 0.9	– 3.0
Spain	1985-89	1.7	6.8	– 0.5	0.6	5.7	– 2.0	– 0.3	– 3.3
	1990-94	1.4	5.9	– 1.6	1.7	8.7	– 6.7	– 1.1	– 3.9
	1995-99	– 1.3	3.1	– 1.9	– 1.2	2.1	– 2.8	*1.4*	*2.3*

Table 7 (cont)

Country	**Period**	**Gap**	**Δpc**	**Δ(Δpc)**	**U*– U**	**Δw**	**Δ(Δw)**	**Δ(Δpc)/Gap**	**Δ(Δw)/(U*–U)**
Sweden	1985-89	1.7	6.2	0.0	0.5	8.9	3.8	0.0	7.6
	1990-94	– 1.9	6.2	– 7.2	– 0.5	6.0	– 4.4	*3.8*	*8.8*
	1995-99	– 1.7	1.6	– 2.2	– 0.6	3.7	– 0.9	*1.3*	*1.5*
Switzerland	1985-90	0.8	1.6	– 0.4	0.0	3.9	0.6	– 0.5	.
	1990-94	– 0.3	4.0	– 4.1	0.0	4.3	– 3.7	*13.5*	.
	1995-99	– 2.4	0.7	– 1.3	– 0.2	1.6	– 0.9	*0.5*	*4.5*

Notation: Gap = average output gap; Δpc = average rate of inflation (private consumption deflator); Δ(Δpc) = change in the rate of inflation during period; U* – U = deviation of actual unemployment from the NAIRU estimates discussed in Note 17; Δw = average rate of wage inflation (compensation per employee, business sector); Δ(Δw) = change in the rate of wage inflation during period; Δ(Δpc)/Gap = change in rate of inflation relative to output gap; and Δ(Δw)/(U* – U) = change in the rate of wage inflation relative to deviation of actual unemployment from estimated NAIRU.

Table 8a

Price equations, 1960-90

Variables	United States		Japan		United Kingdom		Canada	
	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained
Intercept	0.83 (3.0)	– 0.05 (0.4)	0.63 (0.8)	– 0.53 (0.6)	1.37 (2.0)	0.25 (0.7)	1.33 (3.4)	0.38 (1.8)
Gap_{-1}	0.20 (2.7)	0.28 (3.2)	0.47 (1.0)	0.81 (1.4)	0.54 (3.4)	0.63 (3.9)	0.20 (2.2)	0.17 (1.7)
Δpc_{-1}	0.68 (7.6)	0.72^c	0.06 (0.3)	0.42^c	0.55 (3.8)	0.63^c	1.05 (7.7)	1.16^c
Δpc_{-2}	0.14 (1.5)	– 0.28 (2.5)	0.14 (1.0)	0.35 (2.0)	0.34 (5.8)	0.37 (2.5)	– 0.22 (1.7)	– 0.16 (1.1)
$\Delta pm - \Delta pc_{-1}$	0.13 (8.7)	0.13 (6.9)	0.09 (2.4)	0.08 (1.7)	0.30 (3.0)	0.34 (5.5)	0.17 (4.3)	0.19 (4.3)
Δw_{-1}	.	.	0.44 (2.6)	0.23 (1.1)	.	.	.	.
R^2	0.94	0.79	0.85	0.57	0.88	0.68	0.90	0.55
Durbin's h	0.14	0.85	– 0.04	2.21	0.41	0.88	1.58	2.50
Standard error	0.61	0.76	1.91	2.47	1.81	1.91	0.90	1.02
F-test, constraint[1]	.	13.0	.	11.96	.	3.64	.	7.39
ΣErrors[2]	2.2	– 1.3	4.4	– 2.3	1.0	– 4.0	6.4	1.7

Notation: See equations (9)-(10) in the text. [1] The 0.05 (0.01) probability value for not rejecting the constraint is for most countries 4.15 (7.50). [2] Cumulative forecast errors (1991-99), static simulation. [c] Constrained estimate.

Table 8a (cont)

Variables	Australia		Spain		Sweden		Switzerland	
	Unconstrained	**Constrained**	**Unconstrained**	**Constrained**	**Unconstrained**	**Constrained**	**Unconstrained**	**Constrained**
Intercept	2.19 (3.7)	0.26 (0.8)	1.04 (0.8)	0.26 (0.5)	1.03 (1.1)	– 0.01 (0.0)	2.52 (5.4)	0.11 (0.3)
Gap_{-1}	0.23 (1.4)	– 0.03 (0.2)	0.54 (2.5)	0.59 (2.9)	0.30 (2.1)	0.36 (2.6)	0.36 (5.1)	0.10 (1.2)
Δpc_{-1}	1.12 (8.2)	1.30[c]	0.68 (3.6)	0.69[c]	0.67 (3.7)	0.73[c]	0.47 (3.0)	0.97[c]
Δpc_{-2}	– 0.38 (2.9)	– 0.30 (1.9)	0.25 (1.2)	0.31 (1.7)	0.20 (1.1)	0.27 (1.6)	– 0.18 (1.5)	– 0.03 (0.2)
$\Delta pm - \Delta pc_{-1}$	0.18 (4.5)	0.17 (3.5)	0.13 (2.9)	0.13 (2.9)	0.09 (2.0)	0.09 (1.9)	0.12 (3.7)	0.17 (3.6)
Δw_{-1}	.	.	.	.	.	.	.	.
R^2	0.88	0.33	0.80	0.38	0.69	0.30	0.81	0.30
Durbin's h	0.16	2.34	– 0.31	– 0.23	– 1.35	– 0.86	0.91	2.39
Standard error	1.34	1.65	2.46	2.42	1.76	1.77	0.99	1.51
F-test, constraint[1]	.	12.7	.	0.44	.	0.99	.	31.2
ΣErrors[2]	14.0	2.7	8.9	5.1	7.2	1.7	2.2	– 0.1

Notation: See equations (9)-(10) in the text. [1] The 0.05 (0.01) probability value for not rejecting the constraint is for most countries 4.15 (7.50). [2] Cumulative forecast errors (1991-99), static simulation. [c] Constrained estimate.

Table 8b

Price equations, 1960-99

Variables	United States		Japan		United Kingdom		Canada	
	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained
Intercept	0.70 (3.6)	– 0.02 (0.2)	0.37 (0.8)	– 0.40 (0.8)	1.25 (2.5)	0.32 (1.1)	1.11 (3.2)	0.38 (1.9)
Gap_{-1}	0.19 (3.1)	0.27 (3.8)	0.29 (1.4)	0.44 (1.7)	0.59 (4.7)	0.67 (5.1)	0.33 (4.0)	0.28 (3.3)
Δpc_{-1}	0.70 (9.1)	0.73^c	0.10 (0.5)	0.53^c	0.58 (4.7)	0.64^c	0.90 (6.6)	0.99^c
Δpc_{-2}	0.13 (1.6)	– 0.27 (2.9)	0.13 (1.1)	0.30 (2.3)	0.29 (2.3)	0.36 (2.8)	– 0.04 (0.3)	0.01 (0.0)
$\Delta pm - \Delta pc_{-1}$	0.13 (10.7)	0.13 (8.3)	0.10 (4.5)	0.11 (3.7)	0.31 (6.9)	0.31 (6.4)	0.15 (3.8)	0.17 (4.1)
Δw_{-1}	.	.	0.44 (3.4)	0.17 (1.2)	.	.	.	.
R^2	0.95	0.79	0.88	0.63	0.90	0.68	0.90	0.45
Durbin's h	0.37	0.85	– 0.14	2.00	0.42	1.12	0.27	0.95
Standard error	0.54	0.76	1.59	2.06	1.59	1.69	1.01	1.09
F-test, constraint[1]	.	17.3	.	18.00	.	5.07	.	6.06
F-test, stability[2]	0.27	0.21	2.25	1.56	0.23	0.23	1.90	1.51
ΣErrors[3]	1.3	– 0.9	1.8	– 2.6	0.9	– 2.4	2.9	0.2

Notation: see equations (9)-(10). [1] The 0.05 (0.01) probability value for not rejecting the constraint is for most countries 4.15 (7.50). [2] The 0.05 (0.01) probability value for not rejecting parameter stability is for most countries 2.70 (4.05). [3] Cumulative forecast errors (1991-99), static simulations. [C] Constrained estimate.

Table 8b (cont)

Variables	Australia		Spain		Sweden		Switzerland	
	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained
Intercept	1.26 (2.8)	0.16 (0.7)	0.55 (0.7)	0.10 (0.2)	0.97 (1.4)	– 0.05 (0.1)	2.13 (5.5)	0.12 (0.5)
Gap_{-1}	0.08 (0.6)	– 0.07 (0.5)	0.45 (2.9)	0.48 (3.2)	0.33 (2.9)	0.32 (2.8)	0.36 (5.6)	0.10 (1.5)
Δpc_{-1}	1.22 (9.6)	1.32[c]	0.73 (4.6)	0.74[c]	0.65 (4.4)	0.73[c]	0.51 (3.7)	1.00[c]
Δpc_{-2}	– 0.39 (3.1)	– 0.32 (2.4)	0.22 (1.3)	0.26 (1.7)	0.19 (1.3)	0.27 (1.9)	– 0.14 (1.3)	0.00 (0.0)
$\Delta pm - \Delta pc_{-1}$	0.18 (4.7)	0.17 (4.1)	0.13 (3.5)	0.13 (3.5)	0.14 (3.0)	0.14 (3.0)	0.12 (4.0)	0.17 (4.1)
Δw_{-1}	.	.	.	.	.	.	.	.
R^2	0.89	0.35	0.86	0.36	0.73	0.37	0.84	0.31
Durbin's h	0.83	2.00	0.07	0.13	– 1.92	– 1.79	1.63	1.69
Standard error	1.35	1.49	2.11	2.09	1.88	1.92	0.97	1.38
F-test, constraint[1]	.	8.14	.	0.43	.	2.51	.	34.9
F-test stability[2]	1.05	0.32	0.22	0.21	1.51	1.67	0.84	0.38
ΣErrors[3]	7.4	2.4	5.2	3.3	4.8	1.9	1.1	– 0.6

Notation: see equation (9)-(10). [1] The 0.05 (0.01) probability value for not rejecting the constraint is for most countries 4.15 (7.50). [2] The 0.05 (0.01) probability value for not rejecting parameter stability is for most countries 2.70 (4.05). [3] Cumulative forecast errors (1991-99), static simulations. [c] Constrained estimate.

Table 9a
Wage equations, 1960-90

Variables	United States		Japan	United Kingdom		Canada	
	Unconstrained	Constrained	Unconstrained	Unconstrained	Constrained	Unconstrained	Constrained
Intercept	0.17 (0.2)	0.24 (1.0)	7.77 (1.5)	2.77 (4.0)	5.12 (4.3)	4.44 (3.9)	4.44 (3.9)
U	.	.	– 3.1 (1.8)	.	.	– 0.36 (2.3)	– 0.41 (2.6)
U – U*	– 0.81 (4.1)	– 0.80 (4.7)	.	.	.	.	.
Δ(U – U*)	.	.	.	– 2.43 (2.6)	– 3.14 (4.0)	.	.
Δpc	.	.	0.73 (4.9)	.	.	1.05 (4.2)	1.13[c]
Δpc_{-1}	0.29 (2.0)	0.30 (2.0)	.	1.15 (5.5)	1.24[c]	– 0.23 (0.9)	– 0.13 (0.5)
Δpc_{-2}	.	.	.	– 0.37 (1.7)	– 0.24 (1.2)	.	.
Δw_{-1}	0.30 (1.5)	0.29[c]	0.22 (1.9)	.	.	.	.
Δw_{-2}	0.42 (2.1)	0.41 (2.5)	.	.	.	.	.
$\text{Log}\,((^{W}/_{PC})/Q)_{-1}$	.	.	–13.4 (1.3)	.	.	.	.
R^2	0.77	0.47	0.93	0.60	0.40	0.68	0.22
Durbin's h (DW)	0.38	1.21	2.33	(1.92)	(1.84)	(1.71)	(1.53)
Standard error	0.91	0.89	1.67	3.45	3.52	1.72	1.77
F-test, constraint[1]	.	0.01	.	.	2.03	.	2.40
ΣErrors[2]	– 2.2	– 1.9	– 7.5	20.2	9.9	0.8	– 6.6

Notation: See equations (11)-(12) in the text. [1] The 0.05 (0.01) probability values for not rejecting the constraint is for most countries 4.24 (7.77). [2] Cumulative prediction errors (1991-99), static simulations. [C] Constrained estimate.

Table 9a (cont)

Variables	Australia		Spain		Sweden		Switzerland
	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained	Constrained	Unconstrained
Intercept	4.06 (3.5)	4.03 (4.1)	0.72 (0.4)	1.03 (1.1)	0.02 (0.80)	0.58 (1.0)	0.11 (0.1)
U	– 0.48 (2.4)	– 0.48 (2.6)	.	.	.	.	.
U – U*	.	.	– 0.56 (1.8)	– 0.55 (1.8)	– 1.30 (1.9)	– 1.29 (2.0)	– 4.74 (3.9)
Δ(U – U*)	.	.	.	.	– 2.56 (1.5)	– 2.27 (2.1)	.
Δpc	1.82 (7.0)	1.83 (7.3)	0.35 (1.4)	0.32 (1.6)	0.83 (3.4)	0.81 (3.7)	.
Δpc_{-1}	– 0.83 (3.1)	– 0.83^c	.	.	– 0.26 (1.0)	– 0.27 (1.0)	.
Δpc_{-2}	.	.	.	.	.	.	– 0.30 (2.5)
Δw_{-1}	.	.	0.67 (3.3)	0.68^c	0.49 (2.4)	0.46^c	0.53 (4.0)
Δw_{-2}	.	.	.	.	.	.	.
Log $(({}^{W}/_{PC})/Q)_{-1}$	.	.	.	.	.	.	– 16.0 (3.5)
R^2	0.76	0.75	0.79	0.20	0.49	0.56	0.87
Durbin's h (DW)	(2.55)	(2.55)	1.72	(1.76)	– 0.59	(2.08)	0.25
Standard error	2.53	2.46	3.12	3.04	2.24	2.19	1.12
F-test, constraint[1]	.	0.00	.	0.05	.	0.06	.
ΣErrors[2]	– 19.1	– 19.4	11.7	13.5	– 13.1	– 9.9	0.69

Notation: See equations (11)-(12) in the text. [1] The 0.05 (0.01) probability values for not rejecting the constraint is for most countries 4.24 (7.77). [2] Cumulative prediction errors (1991-99), static simulations. [C] Constrained estimate.

Table 9b

Wage equations, 1960-99

Variables	United States		Japan	United Kingdom		Canada	
	Unconstrained	Constrained	Unconstrained	Unconstrained	Constrained	Unconstrained	Constrained
Intercept	0.64 (1.2)	0.37 (1.7)	0.57 (0.2)	3.37 (3.2)	2.50 (4.7)	5.23 (4.7)	4.13 (4.1)
U	.	.	– 0.78 (0.2)	.	.	– 0.37 (3.1)	– 0.34 (2.7)
U – U*	– 0.80 (4.7)	– 0.83 (4.9)	.	.	.	.	.
Δ(U – U*)	.	.	.	– 2.69 (4.1)	– 2.97 (5.0)	.	.
Δpc	.	.	0.80 (6.0)	.	.	0.98 (4.7)	1.08[c]
Δpc_{-1}	0.37 (2.7)	0.36 (2.8)	.	1.20 (6.5)	1.24[c]	– 0.15 (0.7)	– 0.08 (0.4)
Δpc_{-2}	.	.	.	– 0.32 (1.6)	– 0.24 (1.4)	.	.
Δw_{-1}	0.30 (1.8)	0.32[c]	0.32 (3.2)	.	.	.	.
Δw_{-2}	0.29 (1.9)	0.32 (2.3)	.		.		.
$\text{Log}\,((^{W}/_{PC})/Q)_{-1}$	.	.	–24.7 (4.3)		.		.
R^2	0.79	0.43	0.94	0.68	0.45	0.77	0.19
Durbin's h (DW)	0.73	1.24	2.31	(1.77)	(1.77)	(1.80)	1.59
Standard error	0.95	0.94	1.65	3.17	3.17	1.62	1.69
F-test, constraint[1]	.	0.32	.	.	0.92	.	4.02
F-test, stability[2]	5.00	5.78	0.98	1.60	1.05	1.89	2.10
ΣErrors[3]	– 0.4	– 1.3	– 2.9	11.7	7.9	0.4	– 3.7

Notation: See Table 9a. [1] The 0.05 (0.01) probability values for not rejecting the constraint is for most countries 4.15 (7.50). [2] The 0.05 (0.01) probability values for not rejecting parameter stability is for most countries 2.28 (3.21). [3] Cumulative prediction errors (1991-99), static simulations. [c] Constrained estimate.

Table 9b (cont)

Variables	Australia		Spain		Sweden		Switzerland
	Unconstrained	Constrained	Unconstrained	Constrained	Constrained	Unconstrained	Unconstrained
Intercept	4.13 (3.9)	3.32 (4.0)	0.21 (0.2)	0.55 (0.8)	2.55 (2.1)	1.00 (2.0)	0.02 (0.0)
U	– 0.29 (2.2)	0.27 (2.0)	.	.	.	.	.
U – U*	.	.	– 0.61 (2.5)	– 0.59 (2.4)	– 1.16 (2.2)	– 0.92 (1.8)	– 5.36 (4.2)
Δ(U – U*)	.	.	.	.	– 1.02 (1.0)	– 1.71 (2.0)	.
Δpc	1.80 (7.9)	1.87 (8.3)	0.44 (2.1)	0.30 (1.9)	0.64 (3.6)	0.71 (4.0)	.
Δpc_{-1}	– 0.92 (4.1)	– 0.87[c]	.	.	– 0.15 (0.8)	– 0.11 (0.6)	.
Δpc_{-2}	.	.	.	.	.	.	– 0.23 (1.9)
Δw_{-1}	.	.	0.65 (4.0)	0.70[c]	0.31 (1.9)	0.40[c]	0.51 (4.0)
Δw_{-2}	.	.	.	.	.	.	.
Log $((^{W}/_{PC})/Q)_{-1}$	.	.	.	.	.	.	– 15.3 (3.4)
R^2	0.80	0.72	0.86	0.21	0.62	0.44	0.87
Durbin's h (DW)	(2.34)	(2.25)	1.75	(1.72)	– 0.74	(2.13)	– 0.29
Standard error	2.32	2.34	2.90	2.80	2.31	2.35	1.25
F-test, constraint[1]	.	1.55	.	1.00	.	2.05	.
F-test, stability[2]	1.28	1.75	1.87	2.25	1.28	1.59	2.13
ΣErrors[3]	– 4.8	– 8.3	5.5	9.4	2.6	– 2.6	0.04

Notation: See Table 9a. [1] The 0.05 (0.01) probability values for not rejecting the constraint is for most countries 4.15 (7.50). [2] The 0.05 (0.01) probability values for not rejecting parameter stability is for most countries 2.28 (3.21). [3] Cumulative prediction errors (1991-99), static simulations. [c] Constrained estimate.

Chart 1

Errors: consumer prices

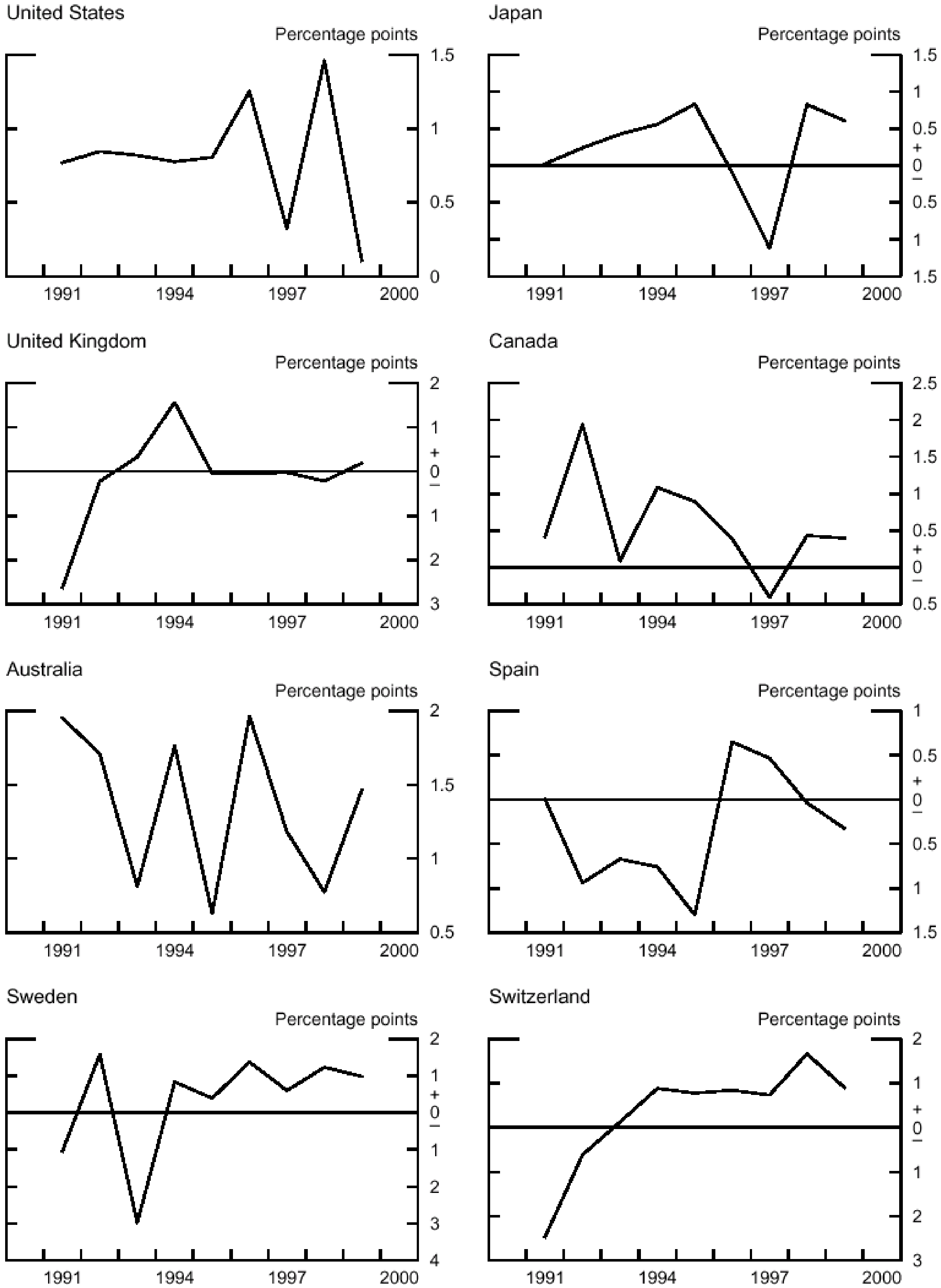
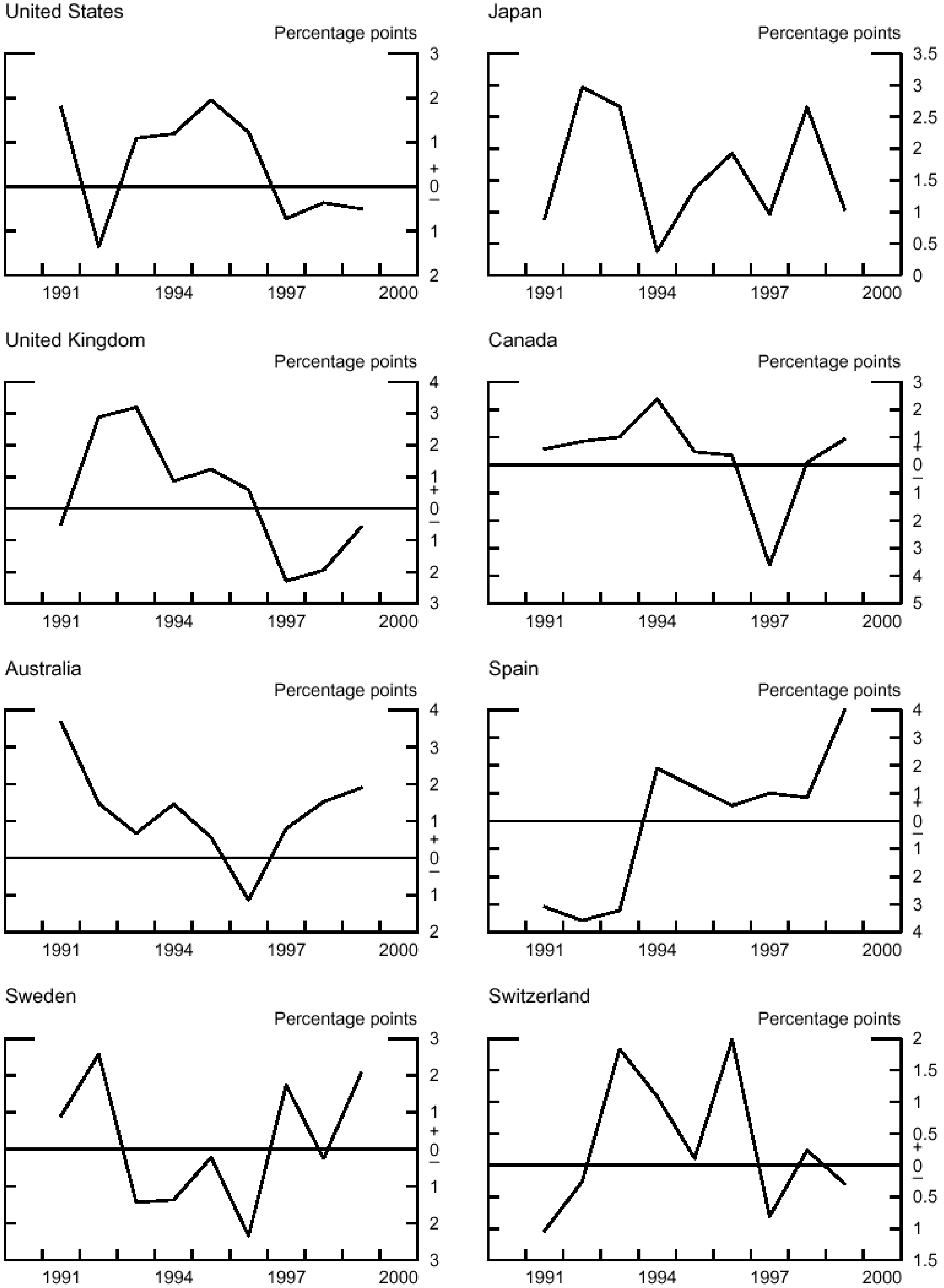


Chart 2

Errors: compensation



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