What explains inflation in high income countries? Evidence from panel data.

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Abstract

Models of monetary policy incorporating imperfectly competitive labour markets predict that inflation increases with the gap between actual real wages and their competitive levels. We use panel data to test for a link between inflation and the labour market characteristics that influence wage gaps. We also test for effects from central bank independence, trade openness and income per capita. Trade union membership rates, employment protection and coordination in wage bargaining are significant. Central bank independence, trade openness and income per capita do not explain temporal variation in inflation directly, but do affect the impact of labour market characteristics on inflation.

KEYWORDS: Inflation, trade union density, labour markets, central bank independence, trade openness.
JEL CLASSIFICATION: E31, J51.

1 Introduction

Explaining differences in inflation across macroeconomic regimes has become an important research topic in monetary economics. Theoretical models of the sort proposed by Barro and Gordon (1983) predict that equilibrium, or time consistent, inflation is a function of those factors determining a policy-maker's degree of inflation aversion. These include the conservativeness of the central bank (Rogoff (1985)) and openness to international trade (Romer (1993), Lane (1997)).

The recent contributions to this literature emphasise the importance of labour market institutions. Cukierman and Lippi (1999), hereafter CL, consider a two-stage game featuring trade unions and a central bank. The key result is that trade unions set real wages in excess of the competitive level and that it is optimal for the central bank to partially accommodate this real wage premium when setting inflation. CL argue that trade union centralisation (the inverse of the number of unions) is the main determinant of the real wage premium and hence inflation, but it is clear that other factors such as the percentage unionisation rate of the workforce, the substitutability of employed and unemployed labour and the degree of coordination between firms and unions could also play a role.

Empirical evidence on the predictions made by these models has been presented in a number of papers published since the start of the 1990s. Cukierman (1992) identifies a robust link between measures of central bank independence (CBI) and inflation. Romer (1993) estimates a negative correlation between trade openness and inflation, and CL provide evidence on the relationship between trade union centralisation and inflation. These findings are obtained either from cross-country regressions or from pooled OECD time series regressions that do not control for country fixed effects. This means that the results may be driven by some time invariant factor that is omitted from the analysis, but which happens to be correlated with the variables of interest.

In this paper we construct a panel dataset for 20 OECD countries covering the period 1961-95. The dataset comprises a range of labour market indicators such as unionisation rates, measures of coordination in wage bargaining and the strictness of employment protection, as well as the variables emphasised in earlier contributions such as central bank independence, trade openness and per capita income. Each variable is time-varying. Using this dataset we investigate the determinants of inflation after controlling for country fixed effects, common time effects and lagged inflation. A second contribution of the paper is that we test for a range of interactions between labour market institutions, CBI, trade openness and income per capita.

The results point to a number of interesting conclusions. First, CBI, trade openness and income per capita are all negatively associated with inflation in regressions that exclude fixed effects, but each effect loses significance when the models control for fixed effects. Second, the percentage unionisation of the workforce exerts a positive and robust effect on inflation. Third, the positive impact of unionisation on inflation is larger when levels of employment protection are above the sample average and smaller when employment protection is below the sample average. Fourth, although CBI, trade openness and per capita income are insignificant when entered as level effects, they do form significant interactions with the unionisation rate. Specifically, the trade union density effect is less strong when income per capita and CBI are above the OECD average, and, in the case of relatively open economies, when the level of coordination between trade unions and firms is above the OECD average. Finally, we show that the impact of oil price shocks on inflation varies with the structure of labour market institutions. These findings are, for the most part, robust to controlling for reverse causation bias and to varying the sample used for estimation.

The remainder of the paper expands upon these points and is structured as follows. Section 2 reviews the theoretical and empirical literature on inflation performance in OECD countries. Section 3 describes the panel dataset that we construct and discusses econometric methodology. Section 4 reports our empirical results and section 5 summarises the paper.

2 Models of inflation performance

The main theoretical results arising from the modern approach to explaining inflation can be illustrated using a stripped down version of the model in CL, which extends the well known work of Barro and Gordon (1983). The CL model consists of a two-stage game between a central bank and trade unions. In the second stage the central bank chooses inflation taking the nominal wages previously set by unions as given. In the first stage each union chooses a nominal wage rate taking the nominal wage rates chosen by all other unions and the subsequent central bank reaction as given. Trade unions derive positive utility from high real wages, but dislike unemployment and inflation, such that their loss function can be written as:

$$\Omega_j = -2w_{r\,j} + Au_j^2 + B\pi^2 \tag{1}$$

where u_j is the rate of unemployment among members of union j, $\pi = p - p_{-1}$ is the rate of inflation (the first difference of the log price level) and A and B are positive parameters. The central bank is averse to volatility in unemployment and inflation about their target values (which are normalised to zero) and thus tries to minimise the following loss function:

$$\Gamma = u^2 + I\pi^2 \tag{2}$$

where I measures the relative inflation aversion of the central bank.

The aggregate labour demand equation in the economy is given by

$$L^d = \alpha(d - \overline{w}_r)L \tag{3}$$

where L^d denotes labour demanded, $\overline{w}_r$ is the log of the average real wage, L is the total supply of labour and α and d are parameters. CL note that this can be reformulated as the following aggregate unemployment equation:

$$u = \frac{L - L_d}{L} = \alpha(\overline{w} - \pi - p_{-1} - w_r^c) \tag{4}$$

where $\overline{w}$ is the average nominal wage, p_{-1}, is the log of the previous period price level and $w_r^c = d - 1/\alpha$ is the market clearing wage at which $u = 0$. Minimising (2) subject to (4) yields the following monetary policy reaction function:

$$\pi = \frac{\alpha^2}{\alpha^2 + I}(\overline{w} - w_r^c - p_{-1}) \tag{5}$$

If (5) is rewritten after splitting the nominal wage into its real and expected price level components ($\overline{w} = \overline{w}_r + Ep$) then the solution for the inflation rate is as follows:

$$\pi = \frac{\alpha^2}{\alpha^2 + I}(\phi + E\pi), \quad \phi = \overline{w}_r - w_r^c \tag{6}$$

which says that the central bank partially accommodates the expected rate of inflation, $E\pi$, and what CL term the 'real wage premium', ϕ, which is the excess of the equilibrium real wage over the competitive (full employment) level. Imposing the rational expectations condition that $\pi = E\pi$ yields the following expression for inflation:

$$\pi = \frac{\alpha^2}{I}\phi \tag{7}$$

Equation (7) illustrates the well known result that inflation exceeds the socially optimal level (of zero) when real wage rigidities prevent employment from reaching the competitive level. The intuition for this is that if private inflation expectations were equal to zero the policy-maker would have an incentive to 'cheat' on those expectations through moving the economy along the short-run unemployment-inflation trade-off in (4) to a point at which there is positive inflation and lower unemployment. This is welfare improving in the short-term because the policy-maker has convex preferences over low inflation and low unemployment and therefore benefits from trading one off against the other. The private sector anticipates this move and therefore sets inflation as in (7), at which point inflation is sufficiently high to deter further episodes of surprise inflation.

The expression in (7) can be used to elucidate a number of arguments that have been put forward in order to explain inflation. These theories can be divided into three groups: those relating to the preferences of the policy-maker, those that focus on labour market institutions and those that look at interactions between preferences and institutions. We summarise these arguments below and indicate in brackets the sign of the predicted effect.

The determinants of inflation preferences

- *The degree of central bank independence (-ve).*

In the models due to Barro and Gordon (1983) and Rogoff (1985) it is implicitly assumed that the real wage premium is constant so that there is a fixed distance between equilibrium levels of employment/output and the socially optimal level. It then follows that inflation will be inversely related to the inflation aversion parameter, I. Assuming that a committee of central bankers is more inflation averse than an elected government, it follows that greater central bank independence (CBI) in the conduct of monetary policy will lead to lower inflation.

- *Openness to international trade (-ve).*

Lane (1997) presents a model in which a wedge between equilibrium and socially optimal employment exists in the non-traded sector of the economy but not the traded sector. In terms of the CL framework this is equivalent to there being a real wage premium in the non-traded sector but not the traded sector (more generally, the ϕ parameter is smaller in the traded sector than the non-traded sector). One rationale for this is that the traded sector is more competitive. In this set-up a policy authority will only have an incentive to implement surprise inflation in the non-traded sector (assuming an initial inflation rate of zero), and as this sector is smaller in a more open economy the overall incentive to launch inflation surprises will be weaker and the equilibrium inflation rate lower.[1]

- *GDP per capita (-ve).*

The inflation aversion parameter, I, may be related to income per capita if low inflation is considered a normal good. This relationship is not set out in a specific model, but is often cited in empirical studies of inflation performance, see Campillo and Miron (1997).

[1] Romer (1993) derives the result that inflation is negatively related to openness using a model in which terms of trade adjustments decrease the net marginal benefit of surprise monetary expansions. The logic of this model is more difficult to illustrate in the framework used here, however.

Labour market institutions

Recent research due to Nunziata (2004) shows that real wages in OECD countries depend on labour market institutions. The solution for inflation in (7) suggests that these institutional variables will also affect inflation. The key labour market characteristics emphasised by Nunziata are as follows:

- *The percentage unionisation rate of the workforce (+ve).*

As unionisation rates increase, the substitutability of union labour and non-union labour decreases, which strengthens the bargaining position of trade unions such that they can elicit a larger real wage premium. This leads to higher equilibrium inflation.

- *The degree of trade union centralisation/coordination (+ve/-ve).*

The degree of centralisation in the trade union sector is the inverse of the number of unions, see CL (1999). This is closely related to the level of coordination between trade unions and firms in the wage bargaining process because greater centralisation facilitates coordination. CL predict a hump shaped relation between centralisation and the real wage premium/inflation. The reason for this is that increasing the average size of a union exerts two effects on union behaviour. First, increased monopoly power in labour supply raises the real wage premium that a union can extract. Second, as unions increase in size the macroeconomic consequences of their wage decisions become more visible. When a small union increases its nominal wage demand there is only an infinitesimal aggregate price change so that the result is a higher real wage without higher inflation, whilst when a large union increases its nominal wage demand the overall price level increases, such that inflation rises and the real wage increase is moderate. This reduces the incentive for trade unions to push for higher real wages. At low levels of trade union centralisation the first effect dominates, meaning that the wage premium rises and inflation increases, whilst after a certain point the second effect dominates and inflation falls. This argument is close to that of Calmfors and Driffill (1988).

- *The strictness of employment protection legislation (+ve).*

An increase in the strictness of employment protection legislation ensures that it is less easy to substitute labour that is currently employed for outside labour, and this will increase the real wage premium that can be earned by those currently in employment.

Interaction effects

The literature points to a large number of interaction effects in the determination of inflation. These can be summarised as follows:

- *Interactions between labour market institutions.*

Inflation may depend on the interaction between labour market institutions. We hypothesise that unionisation rates will exert a larger impact on inflation when employment protection is relatively strict, but will exert a smaller impact on inflation rates when wage bargaining is relatively centralised/coordinated.

- *Interactions between labour market institutions and indicators of inflation aversion.*

Soskice and Iversen (1998) contend that trade unions will be deterred from setting a large real wage premium if central bank independence is high because trade unions anticipate that the central bank will respond to a large real wage premium through implementing a very tough monetary policy. The success of the Bundesbank in 'disciplining' German trade unions is one example of this interaction.

CL highlight a mechanism through which central bank independence may amplify rather than restrict the inflation increasing effect of trade union power. It is shown that the critical level of union centralisation at which the hump reaches a peak depends upon the level of central bank independence. The more independent the central bank the higher the level of union centralisation at which the turning point occurs, i.e. the greater the range of centralisation levels over which inflation will be an increasing function of union centralisation. The intuition for this is that high levels of CBI ensure that inflation is held low when a trade union increases its nominal wage demand. This allows each union greater flexibility in trading off employment and real wages (because higher nominal wages are less likely to be cancelled out by high inflation). To the extent that unions target higher real wages, equilibrium inflation will rise.

2.1 The empirical evidence

A large number of papers provide empirical evidence on the determinants of inflation. The main tools used in these studies are either cross-country regressions or pooled time series regressions that do not control for country fixed effects, i.e. the information used to test the theoretical predictions is mainly based upon differences between countries, rather than changes within countries over time. Examples of these studies include Cukierman (1992), who shows that inflation is negatively related to a measure of the legal independence of central banks and Romer (1993), who shows that average inflation over the period 1973-89 is negatively related to both openness to trade (measured as the share of imports in GDP) and real per capita income, both averaged over the period 1973-89.[2]

CL provide preliminary evidence on the relationship between labour market institutions and inflation. Average inflation rates over 5 year periods are regressed on three dummy variables representing low, intermediate and high levels of trade union centralisation, plus the interactions between the dummies and the Cukierman index of CBI. The results support the theoretical predictions made by CL. A more detailed study is reported in Hall and Franzese (1998), who regress inflation on trade union centralisation, the unionisation rate, per capita income and openness to trade. Regressions are estimated using cross-sectional averages, a stacked time series in which inflation is measured over 10 year intervals, and stacked time series for annual data. The results show that union density, union centralisation and CBI are all statistically significant and correctly signed. The significance of trade openness and per capita income is more marginal, whilst the interaction term suggested by CL is insignificant.

A number of criticisms have been levelled at the empirical literature on inflation performance. First, the estimated relationships could be driven by any time invariant effect that is relevant in setting inflation. For instance, Posen (1993) shows that the cross-sectional correlation between CBI and inflation disappears after controlling for a measure of financial sector opposition to inflation, though the sample considered is smaller than that used by Cukierman (1992). Second,

[2] Sachsida et al (2003) and Gruben and McLeod (2003) control for country effects in evaluating inflation performance, but focus on just one explanatory variable, trade openness.

cross-sectional studies can prove sensitive to the period over which average inflation rates are measured, e.g. Bleaney (1999) shows that the strength of the evidence for a relationship between openness and inflation depends on whether the variables are measured over the 1970s, 1980s or 1990s. Third, many studies fail to control for the full range of variables suggested by economic theory, or do not investigate the possible interactions between them.

3 A panel model for inflation performance in the OECD

The analysis in this paper uses panel data on inflation regimes in 20 countries over the period 1961-95.[3] The time variation in this dataset allows us to test whether or not labour market institutions and indicators of inflation aversion explain the changes in inflation that have occurred over time, as well as the cross-country differences in inflation. In the econometric analysis the full range of interaction effects suggested by economic theory are explored. We now describe both the dataset and the econometric approach.

3.1 Data

The variables included in the dataset are defined below. The methods and sources used in constructing the data are described in the appendix.

• *INFLATION* is the annual rate of consumer price inflation measured as a decimal (1% inflation is recorded as .01 in the dataset).

• *CBI* is an updated version of Cukierman's (1992) index of the legal independence of central banks, provided by van Lelyveld (2000). The range for this index is $0 - 1$, where 1 indicates maximum possible independence. The updated index due to van Lelyveld shows greater time variation than the original Cukierman index.

• *OPEN* is the nominal value of imports plus exports divided by the value of nominal GDP.

• *GDP* is the natural log of real GDP per capita, measured in terms of trade adjusted US$.

• *TU* is the union membership rate for employees, often referred to as trade union density. The feasible range for this variable is $0 - 1$.[4]

• *COORD* is a direct measure of the degree of consensus between actors in collective bargaining. The index lies in the range $1 - 3$, where 3 denotes the highest level of coordination. This variable is closely related to the measure of centralization in union bargaining used by CL. One advantage of *COORD* over the centralisation index is that it is available for a longer period of time and at a higher frequency. Furthermore, as it aims to capture the degree of consensus between firms and unions in the labour market, it should provide a better measure of the extent to which trade unions are likely to moderate wage claims in order to prevent episodes of high inflation.

[3] The countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States

[4] The labour market variables for Germany refer to West Germany only. Note, however, that German unification affects only the last 5 years of the sample, 1991-95.

• EP measures the strictness of employment protection legislation. It takes values in the range $0-2$, where 2 is the highest possible level of employment protection.

In addition to the variables described in the text, Nickell et al (2002) provide data on the generosity and duration of unemployment benefits and on average tax rates. We do not consider variables relating to the benefits system because they may be endogenous to inflation, e.g. high inflation may erode the real value of benefits in the absence of full indexation. We have estimated models that control for tax rates but the effects turned out insignificant and are not reported.

In Figure 1 we plot average inflation in the 20 countries in our sample against the corresponding averages for *TU*, *CBI*, *OPEN* and *GDP* (the plots have been adjusted so that they have comparable means and ranges and therefore we do not report units for the vertical axis). An important feature of the series for *TU* is that it appears to be able to account for the upturn in inflation between the early 1960s and late 1970s as well as the reduction in inflation from the early 1980s onwards.

Figure 0 - see end of document.

3.2 Econometric methodology

In order to examine the determinants of inflation performance we estimate regressions of the following form:

$$INF_{it} = \gamma_0 + \boldsymbol{\gamma}_1'\mathbf{x}_{1,it} + \boldsymbol{\gamma}_2'\mathbf{x}_{2,it} + \boldsymbol{\gamma}_3'\mathbf{h}_{it} + \mu_i + \lambda_t + \varepsilon_{it} \tag{8}$$

In this notation i refers to a country and t refers to a 5 year period. Thus, when $t = 1$ in the panel, each of the observations is an average over the five years $1961-65$. The dependent variable, INF, is defined as ln $(1 + INFLATION)$. This transformation is used because $\ln INFLATION$ can exaggerate the effect of very low inflation countries. The vector $\mathbf{x}_1$ comprises determinants of a country's likely degree of inflation aversion, $\mathbf{x}_2$ is a vector of labour market institutions indicators, $\mathbf{h}$ is a vector of interactions among the first two sets of variables, μ_i a country fixed effect, λ_t a time dummy and ε_{it} the error term.

Inflation is measured over 5 year periods because annual inflation rates will depend on macroeconomic shocks occurring in that year and need not reflect the target inflation rate of the policy authority given the labour market structure that it faces (Ireland (1999) notes that positive theories of inflation do not explain high frequency movements in inflation). This approach to testing models of inflation performance is well established in the literature, see for instance Hall and Franzese (1998) and Gruben and McLeod (2003).

Note that when testing whether or not a variable x_{m2} interacts with a variable x_{n2} in setting inflation, we use the terms $\gamma_{m2} \cdot x_{m2} + \gamma_{m3} x_{m3} \cdot x_{n3}$ in the regression, and define x_{n2} so that it has a zero mean. This ensures that the coefficient on the level of x_{m2} can be interpreted as the coefficient of the "average" country, i.e. the country characterized by the average level of x_{m2}. In the results section a variable preceded by Z indicates that the variable is in zero mean form.

Equation (8) controls for country fixed effects (μ_i) and common time effects (λ_t). The model is estimated using feasible GLS, allowing for groupwise heteroskedasticity and an AR(1) structure in the disturbances (a common error autocorrelation parameter is assumed for the 20 countries in the panel).

4 Empirical results

The first models that we estimate control for time dummies but not fixed effects and therefore emphasise the cross-country variation in the data. In column (1) of Table 1 the explanatory variables are *CBI*, *OPEN* multiplied by a dummy that is equal to 0 up until 1981-85 and equal to 1 after that time (for each country), and *GDP*. Each of the variables is negatively signed and significant, in line with the results from past research. In common with Gruben and McLeod (2003) we find that openness is significant only in the post-1985 period. Entering openness in column (1) without scaling by the dummy yields an insignificant coefficient estimate, results not reported here.

In column (2) we consider the CL hypothesis that inflation is hump shaped in trade union centralisation provided that *CBI* is relatively low. *COORD* is closely related to the centralisation index used by CL and is used in testing the hump shape hypothesis. In column (2) the explanatory variables are *ZCOORD*, *ZCOORD*ZCOORD* and the interaction of those two variables with *ZCBI*. The CL hypothesis implies that *ZCOORD*ZCOORD* should be negatively signed, *ZCOORD*ZCBI* positively signed and *ZCOORD*ZCOORD*ZCBI* negatively signed (this combination of signs implies that the turning point for a graph of inflation against *ZCOORD* occurs further to the right). There is some support for these predictions, but only two of the four parameters are significant at the 5% level. Further, when *CBI*, *OPEN8695* and *GDP* are added in column (3) the interaction terms become very insignificant.[5]

In column (4) we add fixed effects to the column (1) specification. The variables *CBI* and *OPEN* are much less significant, indicating that whilst these variables explain cross-country differences in inflation there is no evidence that they account for changes in inflation within countries. In column (5) we add fixed effects to the column (3) specification. The results indicate some support for the GDP effect and the CL hypothesis.

Table 1 - see end of document.

Labour market institutions In Table 2 we consider a range of labour market variables. The model in column (1) controls for time dummies, fixed effects and trade union density. The coefficient on the unionisation rate is positively signed and highly significant, supporting the view that monopoly power in the labour market increases equilibrium inflation. In column (2) we add EP and $COORD$, but neither term is significant at the 5% level. In column (3) we interact TU with $ZCOORD$ and ZEP. It appears that higher levels of coordination between firms and unions moderate the inflation increasing effect of more heavily unionised labour markets, while higher levels of employment protection make the effect stronger, although the latter interaction

[5] Guzzo and Velasco (1999) show that inflation is hump shaped in CBI and that the shape of the hump depends on coordination in wage bargaining. We briefly investigated this hypothesis but did not find any supporting evidence - allowing for non-linearities in the relationship does not revive a CBI effect.

is insignificant at this stage. Column (4) shows that conditioning on labour market institutions indicators does not restore the significance of CBI and $OPEN$ (this does not change if we replace $OPEN$ with $OPEN8695$).[6]

In column (5) trade union density is interacted with zero mean versions of CBI, GDP and $OPEN$, and in column (6) three triple interaction terms are added to the model. The interaction between union density and income per capita is significant, but the other coefficients are less well determined. In column (7) we report a tested down version of equation (6), obtained through deleting the least significant term, re-estimating the model and then repeating the process until each term is significant at the 5% level. The results confirm the importance of trade union density. The effect is stronger when levels of employment protection are above the sample average and weaker when employment protection is below average. These findings are consistent with the results in Nunziata (2004), who shows that high levels of union density and employment protection increase real wages relative to productivity. This implies higher unit costs, which may subsequently increase inflation.

Central bank independence above the OECD average reduces the impact of trade union density on inflation. One interpretation is that trade unions are less likely to submit high wage demands when the policy authority is tough because they anticipate a significant tightening of monetary policy, which would drive unemployment to a level that unions cannot tolerate. One example of this is the interaction between German trade unions and the Bundesbank, see Soskice and Iversen (1998).

In column (7) an above average level of income per capita reduces the impact of union density on inflation. This may reflect a tendency for unions to moderate pay claims when living standards are relatively high. It is important to note that as per capita income is a trended variable the effect of union density is stronger at the start of the sample than at the end. If we replace the GDP per capita term with GDP per capita relative to the OECD average at that time, which is not a trended variable, we obtain a coefficient estimate for *TU*ZGDP* of $-.065$ (absolute *t-ratio* is 1.64) and a coefficient on *TU* of .036 (absolute *t-ratio* is 1.71), and all of the other regressors remain significant at the 5% level. If the GDP interaction is dropped from the model then all of the regressors, including *TU*, are significant at the 5% level and the coefficient on *TU* is .040. One explanation for the robustness of the results after deleting a heavily trended interaction term is that the time dummies, λ_t, control for its effects.

The column (7) results also indicate that a relatively high level of coordination between firms and workers moderates the inflationary impact of union density, but only in countries in which $OPEN$ exceeds the OECD average. This lends support to the hypothesis due to Rama (1994), that union behaviour is less aggressive in open economy settings because international competition raises the elasticity of labour demand and hence the unemployment cost of high real wages.[7] The interaction *TU*ZCOORD* was deleted during the model reduction process.

[6] Daniels, Nourzad and VanHoose (2003) argue that openness reduces inflation by more in nations with less centralised wage bargaining. The implied interaction term is not significant in models that control for fixed effects and time dummies (results not reported here).

[7] It will be noted from equation (7) that a higher elasticity of labour demand can increase the pass-through from the real wage premium to inflation. The negative coefficient on the three-way interaction term suggests that this effect is dominated by the Rama effect, which says that trade unions increase the wage premium by less

If it is added to model (7) the estimated coefficient is negative, but the absolute *t-ratio* is just 0.91. The coefficient on the three-way interaction *TU*ZCOORD*ZOPEN* remains significant and has an absolute t-ratio of 2.47. Only if the latter term is excluded does *TU*ZCOORD* become significant at the 5% level (absolute *t-ratio* of 2.49).

If the four variables suggested by the CL hump-shape hypothesis, *ZCOORD, ZCOORD*ZCBI, ZCOORD*ZCOORD* and *ZCOORD*ZCOORD*ZCBI,* are added to column (7), each of them is insignificant.[8] One explanation is that unionisation rates correlate with other forms of trade union power, including centralisation. This is suggested in Chou (2000) and also in Daniels, Nourzad and VanHoose (2003), who estimate a hump shape relation between union density and inflation, rather than union centralisation/coordination and inflation. Testing for this hump shape through adding the square of trade union density to model (7) in Table 2 yields a *t-ratio* that is marginally insignificant at the 5% level (all other variables remain significant). This suggests that it is not large trade unions per se that cause wage pressures to be moderated, but rather that large trade unions are found in countries in which central bank independence, trade openness, per capita income and union-firm coordination are also relatively high, and that these factors limit the positive effect of union density on inflation.

Oil price shocks We now investigate the effects of oil price shocks. The term *OIL* measures the percentage change in the US$ spot price of a barrel of West Texas Intermediate. As this variable takes the same values for each country it cannot be used in levels form, but it can be added as an interaction term. In column (8) we add the terms *OIL*ZCOORD* and *OIL*ZCOORD*ZOPEN* to the column (7) specification. Both effects are negatively signed and significant. The first interaction suggests that greater coordination between firms and unions reduces the impact of oil price shocks on inflation, possibly because corporatist arrangements lead to movements in the labour share that accommodate adverse oil price shocks, see Bruno and Sachs (1985), Chou (2001) and Nielsen and Bowdler (2003). The second interaction is included because model (7) suggests that the effects of trade union coordination tend to be stronger in more open economies.[9]

In column (9) we delete the time dummies from (8) and add *OIL*. The additional term is highly significant. The other effects are generally robust, although the interaction between union density and GDP loses significance following the deletion of the time dummies. The robustness of the union density effect in model (9) casts some light on the direction of causation in the estimated regressions. One potential objection to the results in columns (1)-(8) is that the union density effect is endogenous in the sense that oil price shocks raise inflation and also cause workers to join trade unions in an effort to protect real incomes. The fact that union density remains significant after controlling for oil price effects suggests that the extent of this reverse causation bias is quite limited (in section 4.2 we address other sources of endogeneity).

We also considered the hypothesis due to Walsh (1997) that relatively high levels of CBI

when labour demand is relatively elastic.

[8] This is consistent with the evidence in Hall and Franzese (1998). The theoretical underpinnings for the CL hypothesis have been questioned by Fracasso and Ozkan (2004) and Ciccarone and Marchetti (2002).

[9] Interactions between oil price inflation and union density and oil price inflation and employment protection were investigated, but the estimated effects proved to be insignificant.

moderate the impact of oil price inflation, but were unable to find any supporting evidence (results not reported).

Stability across macroeconomic regimes Column (10) in Table 2 considers the stability of the preferred specification (model (8)). It is often argued that the Bretton Woods fixed exchange rate system influenced the behaviour of inflation during the period through 1972. In order to examine this possibility we estimate regression (8) for the sub-sample of observations beginning in 1976 – 1980. The results show that most of the estimated effects are robust, although some of them are less well estimated. The CBI interaction loses significance. This indicates that in the full sample the role of CBI is mainly due to events in the 1960s and early 1970s. An inspection of the data shows that CBI declined somewhat in Austria, France and the UK during this period, possibly reflecting the way in which governments instructed their central banks to manage currency flows in supporting exchange rate targets (consider the experience of the United Kingdom in 1967). These reductions in CBI appear to have amplified the inflation increasing effect of rising trade union density during the 1960s and 1970s, and therefore explain a large part of the negative point estimate on $TU*ZCBI$ in column (8).[10]

Table 2 - see end of document.

The quantitative significance of the results A hypothetical country that is exactly at the OECD average in terms of EP, GDP, CBI, $COORD$ and $OPEN$ experiences an increase in the annual inflation rate of 0.68 percentage points following a 10 percentage point increase in union density (using the column (8) results). This effect is very small, but it is important to bear two points in mind. First, this is the effect observed after controlling for the common global trends in inflation and union density and is therefore a lower bound on the effect in which we are interested (leaving aside the issue of estimation uncertainty). An upper bound may be obtained by deleting the time dummies from column (8). This yields a coefficient on union density of 0.217, implying that a 10 percentage point increase in unionisation raises equilibrium inflation by approximately 2.2 percentage points. Second, changes in inflation could be larger still given a particular institutional configuration, e.g. high levels of employment protection and low levels of CBI.

The next question that we address is whether or not the marginal effect of union density is always positive, given the many negatively signed interaction terms that enter the models. In Table 3 we list the coefficient estimates from column (8) in Table 2 for terms in union density, the standard deviations of the zero mean parts of those terms and the products of the coefficient estimates and the standard deviations. The main point is that after controlling for differences in the variability of the interaction terms, it is clear that there is very little chance that the total derivative of inflation with respect to union density will turn negative for any of the countries

[10] During the sub-period commencing 1976-80 many European countries adopted fixed exchange rate regimes as part of the European Monetary System (EMS). Excluding all of these countries would lead to a very small sample and is not an option that we pursue here. Note that if EMS membership affects only the level of inflation then the country dummies will control for its effects.

in the sample (only in the case of Switzerland is the total effect of union density approximately zero).

Table 3 - see end of document.

4.1 Robustness and sensitivity

Figures 1 to 7 show plots of recursive coefficient estimates for the preferred model (Table 2, column 8) obtained by deleting one country at a time from the sample. The plots confirm the stability of the coefficients and suggest that the results are not driven by outlying observations observed in one particular country. The *t-ratio* for $TU * ZCBI$ when Austria is excluded from the sample is 1.44. Unionisation rates in Austria fell during the 1960s and early 1970s at a time of rising inflation. The fact that the independence of Austria's central bank was being eroded at the time helps to explain rising inflation despite falling unionisation rates. This episode seems to be important in driving the full sample estimate of the coefficient on $TU * ZCBI$.

The inclusion of Sweden in the sample is important in obtaining a precise estimate of the coefficient on the employment protection interaction, although the point estimate is robust to dropping observations for Sweden. The coefficient on the GDP interaction is less significant when countries at the extremes of the income distribution are dropped, e.g. Denmark, Japan and Portugal. On the whole, however, the estimated model is quite stable.

Figures 1-7 - see end of document.

Another check that we carried out involved dropping one variable at a time from the specification in Table 2, column (8). In most cases the variables that were left in the model remained significant, suggesting that the core results are not dependent on over-fitting bias (results not reported).

We now consider whether or not the main results (column (8) in Table 2) are dependent on the econometric approach that we have followed. In Table 4 we report the results obtained from the following experiments:

• In column (1) we add an extra set of time observations to the model. These are formed from three year averages for $1996-98$ (most of the series end in 1998). For Canada and Spain the final observation for union density is a $1996-97$ average, while the observation for Belgium is missing. Also, a measure of openness for Australia is unavailable for $1996-98$. The CBI variable is unavailable post-1995. We assume that the observations for $1996-98$ are equal to those for $1991-95$. This is a strong assumption, but it at least permits estimation of the model using a larger sample.

• In column (2) we present a fixed effects plus time dummies specification estimated by OLS and using White (1980) robust standard errors, rather than the feasible GLS method described in section 3.

• In column (3) we generalise the static equations to include a lagged inflation term. If the effects that we estimate in Table 2 are spurious then we would not expect them to be robust to the inclusion of a lagged dependent variable. The dynamic model in (3) is fitted by OLS and does not control for fixed effects. The results obtained from such an estimation are inconsistent. We present them here because it is known that they provide an upper bound on the true value of the autoregressive parameter and can be used as a basis for comparisons with consistent estimates.

• In column (4) we add fixed effects to the column (3) specification and again estimate by OLS. The autoregressive parameter estimated for this model is biased downwards and will represent a lower bound on the true value, see Blundell et al (2000).

• In column (5) we estimate the dynamic specification using the system generalised method of moments estimator (GMM-SYS) due to Blundell and Bond (1998). The method involves estimating the dynamic model in first differences and instrumenting all terms using lagged levels, and estimating the relation in levels using lagged first differences as instruments. The final estimates are weighted averages of those from the two halves of the system. In choosing the set of instruments to be used we treat the contemporaneous levels of each of the independent variables as potentially endogenous and therefore use lagged levels dated $t-2$ (and earlier) as instruments for the equation in first differences, and lagged differences dated $t-1$ (and earlier) as instruments for the equation in levels.[11] The estimator deals with endogeneity concerns as well as the presence of dynamics. The estimator is consistent as the cross-sectional dimension of the panel goes to infinity. As we have just 20 countries in this application the results should be treated with some caution.

The coefficient on union density falls by one half when the sample is extended to 1998 and the effect is actually insignificant at the 5% level. The lack of significance is mainly due to the loss of independent information in the estimation. The time dummy introduced for the $1996-98$ observations is highly insignificant ($p-value=.76$) and when it is deleted the union density term becomes significant at the 5% level. Nevertheless, the results suggest that changes in inflation performance between $1991-95$ and $1996-98$ are less highly correlated with changes in trade union density (though bear in mind the assumptions made in constructing the data).

The results based on OLS estimation are in line with those obtained by feasible GLS. Columns (3) and (4) provide the OLS and within groups estimates of the dynamic panel model. We do not assign any particular interpretation to these results in view of the fact that the estimation is biased, though it is worth noting that the signs and significance of most of the variables are preserved, albeit with large changes in the coefficient estimates in some cases (especially after taking into account the multiplier from the lagged dependent variable).

In column (5) we present the GMM-SYS estimates. The residual diagnostics show that the application of the GMM estimator is valid. The Sargan test is based on the null hypothesis that the over-identifying moment conditions are satisfied. Non-rejection of this hypothesis implies that the corresponding elements of the instrument set are valid. The AR(1) and AR(2) tests indicate that the model residuals are negatively correlated at the first order but uncorrelated at the second order. This is consistent with the assumption that the errors in the levels inflation equation are serially uncorrelated (these errors form a first order moving average in

[11] In the first difference equation there is a moving average error structure, such that the first valid instruments are available at $t-2$ rather than $t-1$, even though it is only the t dated terms that are treated as endogenous.

the differenced equation so that AR(1) effects are expected (the standard errors take this into account)).

The results of the GMM estimation show that most of the effects estimated previously are robust to controlling for the lagged dependent variable and potential endogeneity biases. The estimated coefficient on trade union density is very small at .027, but this rises to .05 after multiplying by $\frac{1}{1-.463}$ in order to obtain the static solution to the model. This long-run effect is not too far from the .068 estimated by feasible GLS, and is significant at the 5% level using the standard errors for the static solution. The term $TU * ZCBI$ loses significance when the model is fitted using GMM-SYS and is actually incorrectly signed. The reason appears to be that lagged values of $TU * ZCBI$ are not good instruments because changes in CBI tend to occur infrequently and suddenly rather than in a serially correlated fashion, so that there is only a weak correlation between levels and first differences of the series. The other interaction terms are all significant in the GMM-SYS estimate of the model. The static coefficients multiplying these terms are somewhat larger in absolute value than those reported in column (8) of Table 2, especially in the case of the oil price terms.

Table 4 - see end of document.

4.2 Further results

In this sub-section we consider variables for which we were able to collect data for only a subset of the 20 countries included in the full sample. The variable *LEFT* measures the proportion of cabinet seats filled by left wing parties, while *RIGHT* measures the proportion taken by right wing parties (the two variables are not collinear because some cabinet seats are taken by centrist parties). Data are not available for Portugal and Spain.[12] *FOI* is Posen's (1993) measure of financial sector opposition to inflation, *FOI*, which is available for 16 countries. It is not time-varying and is therefore interacted with time-varying regressors rather than being used as a separate levels term. *HOME* is Oswald's (1996) measure of the proportion of households that are owner occupiers. Oswald suggests that this variable is negatively related to geographical labour mobility because home ownership increases the costs of worker relocation. The data are collected at 10 year intervals and we assume that an observation for, say, 1960 applies for the period $1960 - 69$. We then form the five year averages for the periods $1961 - 65$ through $1991 - 95$ that we require. This ensures that the data are smooth within groups. The series is not available for Portugal.

In Table 5 we present the results obtained using these additional variables.[13] In column (1) we present the core specification from column (8) in Table 2 augmented with the levels of *LEFT* and *RIGHT* and interactions between union density and zero mean versions of *LEFT* and *RIGHT*. The results do not support the idea that left wing governments increase inflation relative to right wing governments, either directly, or through interactions with the unionisation rate. In column (2) we find strong support for the hypothesis that higher levels of financial

[12] The final observation for each country is a 1991-94 average, not a 1991-95 average.

[13] The means of the interaction terms in labour market institutions, openness, CBI and GDP per head are not quite zero in these regressions because of the reduction in the sample size.

sector opposition to inflation reduce the impact of union density on inflation. The interaction between union density and *CBI* remains significant in column (2), suggesting that conservative financial sector institutions not only prevent trade unions raising the inflation rate through lobbying the central bank, but also through other means such as lobbying governments not to accommodate inflationary pay claims. The interaction terms in employment protection and per capita income lose significance in (2), but it is not clear whether this is due to the change in the sample size or the inclusion of the additional regressor. Finally, in column (3) it is shown that high levels of home ownership increase the impact of union density on inflation, but the effect is not significant at the 5% level.

Table 5 - see end of document.

5 Summary

This paper has analysed the determinants of inflation performance in the OECD using panel data models that control for fixed effects and time dummies. The main findings are five-fold. First, higher unionisation rates are associated with higher inflation. Second, the impact of unionisation rates on inflation is larger when employment legislation is strict relative to the OECD average, possibly because the bargaining power of unions is increased so that they can extract larger real wage premia. Third, the impact of unionisation rates on inflation is smaller when wage bargaining is highly coordinated, though only in countries that are more open to international trade is the average OECD country. This is consistent with the hypothesis due to Rama (1994) that large trade unions are more likely to moderate wage demands in open economies because foreign competition increases the wage rate elasticity of labour demand. Fourth, after controlling for fixed effects, CBI, trade openness and GDP per capita do not exert significant effects on inflation. Fifth, these variables do play a role in conditioning the impact of unionisation rates on inflation. Increased CBI may cause trade unions to be less aggressive in wage negotiations because unions anticipate that high wage claims will elicit interest rate hikes. Increased GDP per capita may reduce the impact of trade union density on inflation because unions are less militant when living standards are already at a relatively high level.

In extensions of our core results we showed that the main findings are generally robust to varying the sample, both cross-sectionally and along the time dimension. Similarly, most of the estimated effects proved to be robust to estimating a dynamic version of the model using a GMM technique (the interaction between union density and central bank independence proved least robust).

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Appendix: Data Sources

INFLATION. The data were extracted from the OECD annual national accounts, except in the case of Denmark and the Netherlands, for which the source is the *International Financial Statistics* database maintained by the International Monetary Fund.

CBI. The index is obtained by aggregating indicator variables that describe the conditions under which central banks extend loans to the government, the terms of reference for central bank governors and other relevant aspects of central bank operations. The index is due to Cukierman (1992) and refers to the period $1950-89$. The version that we use in this paper is updated to describe changes in central bank independence up to 1994, and is described in van Lelyveld (2000). Updated measures of central bank independence are not provided for Portugal. We use the Cukierman (1992) data for Portugal.

OPEN. The data were extracted from Penn World Tables. German data for the pre-unification period are not available via the Penn World Tables. We obtained the relevant data from the *International Financial Statistics* database and then spliced this to the 1990s data for Germany in order to obtain a consistent series.

GDP. The source for these data is the Penn World Tables. In the case of Germany we take *International Financial Statistics* data on nominal German GDP in DM and convert it to US$ using a centred 11 year moving average of the actual $-DM exchange rate. This series is then divided by the US price level and German population to give real per capita GDP in US$. Finally, this series is spliced to the Penn series for 1990-98 to give the data that we use for German GDP.

EP. Blanchard and Wolfers (2000) provide an index of employment protection at the 5 year frequency. We use an interpolated version of this series, readjusted in mean.

TU. This is the ratio of employed union members to total employees. For European countries other than Sweden the source is Ebbinghaus and Visser (2000). For the other countries the sources are Visser (1996) and Huber et al (1997). The latter series are updated by Nunziata (2003).

COORD. This variable is obtained by interpolating OECD data on bargaining coordination.

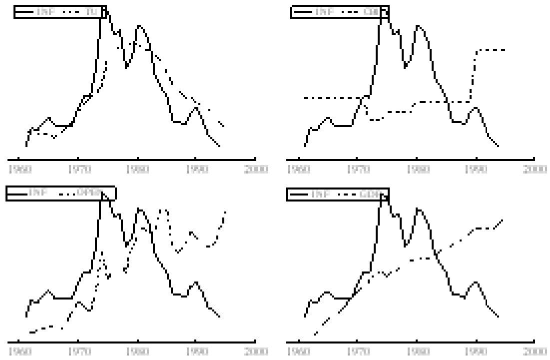

Figure 0: Cross-country averages for INF, TU, CBI, OPEN and GDP, 1961-95.

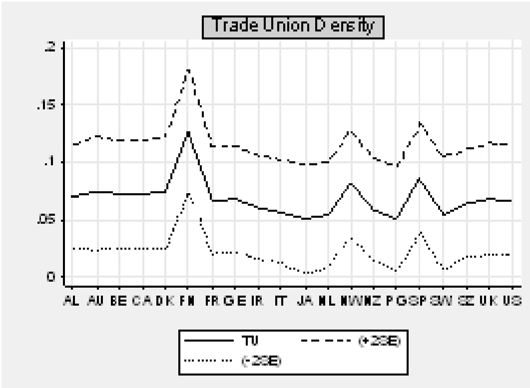

Figure 1: Stability of coefficients in preferred model: TU

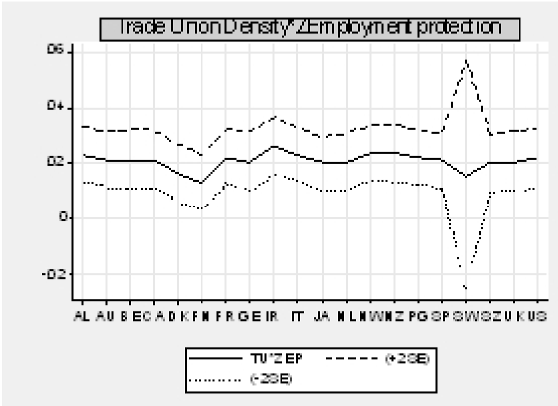

Figure 2: Stability of coefficients in preferred model: TU*ZEP

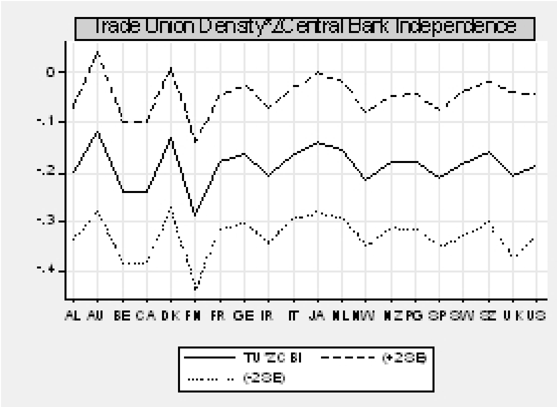

Figure 3: Stability of coefficients in preferred model: TU*ZCBI

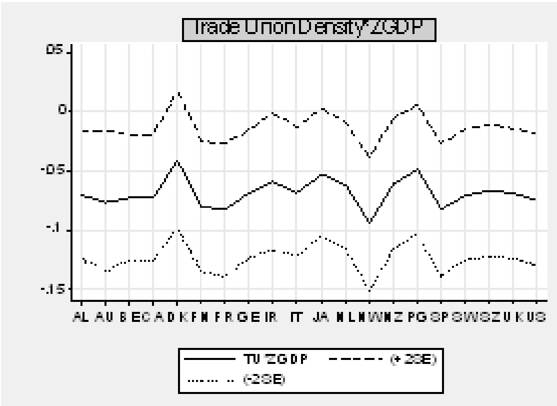

Figure 4: Stability of coefficients in preferred model: TU*ZGDP

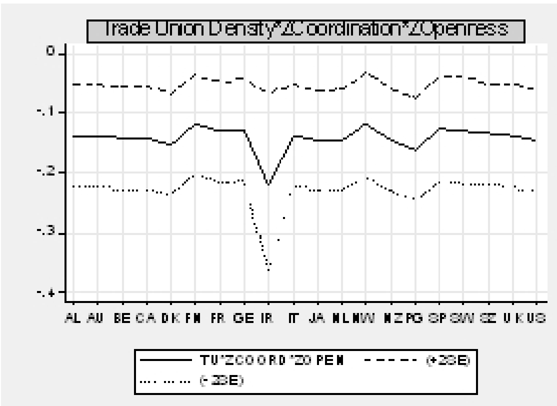

Figure 5: Stability of coefficients in preferred model: TU*ZCOORD*ZOPEN

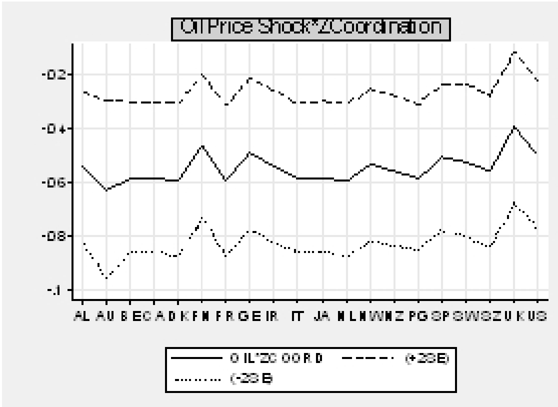

Figure 6: Stability of coefficients in preferred model: OIL*ZCOORD

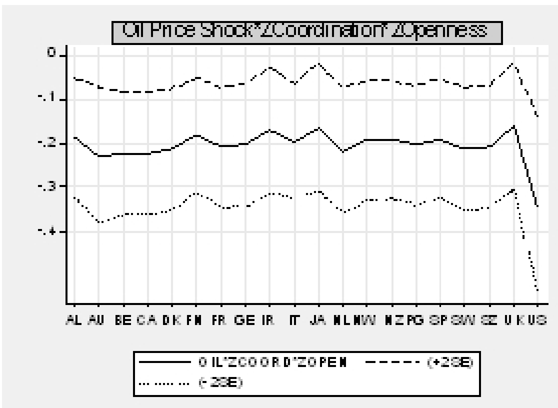

Figure 7: Stability of coefficients in preferred model: OIL*ZCOORD*ZOPEN

Table 1

Basic determinants of inflation in the OECD

Regression	1	2	3	4	5
Constant	.366 (4.78)	.042 (11.16)	.361 (3.83)	.342 (1.94)	.395 (2.33)
CBI	-.034 (2.73)		-.016 (.72)	-019 (.53)	-.007 (.15)
OPEN8695	-.025 (2.39)		-.022 (2.01)	-.012 (1.22)	-.006 (.65)
GDP	-.035 (4.13)		-.034 (3.29)	-.033 (1.71)	-.038 (2.10)
ZCOORD		-.008 (2.28)	-.009 (2.70)		-.001 (.13)
*ZCOORD*ZCOORD*		-.010 (1.64)	-.010 (1.89)		-.022 (1.69)
*ZCOORD*ZCBI*		.044 (1.76)	.017 (.62)		.192 (2.22)
*ZCOORD*ZCOORD*ZCBI*		-.074 (2.20)	-.025 (.66)		.050 (.37)
Fixed effects	No	No	No	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes
No. of countries	20	20	20	20	20
No. of observations	139	136	136	139	136
Root mean square error	0.0244	0.02617	0.0229	0.02155	0.01831
AR(1) error parameter	0.3	0.34	0.27	0.12	0.01

Panel regressions for five-year average inflation rates in 20 countries over the period 1961-95 (inflation measured as a decimal, .01 approximately corresponds to 1%). Each model is estimated by feasible generalised least squares allowing for heteroscedastic AR(1) errors. Figures reported are coefficient estimates, absolute t-ratios in parentheses.

Table 2

Labour market institutions and inflation performance in the OECD

Regression	1	2	3	4	5
Constant	.005 (.40)	.042 (1.54)	.022 (1.61)	.344 (1.96)	-.219 (1.13)
TU	.074 (3.25)	.062 (2.62)	.048 (2.06)	.043 (1.89)	.104 (4.29)
EP		.003 (.70)			
COORD		-.015 (1.52)			
*TU*ZCOORD*			-.047 (2.65)	-.030 (1.57)	-.041 (2.39)
*TU*ZEP*			.007 (.98)	.005 (.68)	.016 (2.49)
CBI				-.027 (.79)	-.029 (.53)
GDP				-.034 (1.78)	.025 (1.18)
OPEN				-006 (.29)	-.041 (.98)
*TU*ZCBI*					-.126 (1.06)
*TU*ZGDP*					-.154 (5.16)
*TU*ZOPEN*					.068 (.88)
Fixed effects	Yes	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes
No. of countries	20	20	20	20	20
No. of observations	137	137	137	136	136
Root mean square error	0.0281	0.01883	0.01853	0.01811	0.01758
AR(1) error parameter	0.17	0.15	0.14	0.06	-0.01

Table 2 continued

Regression	6	7	8	9	10
Constant	-.235 (1.16)	-.006 (.40)	.002 (.15)	-.014 (.84)	.074 (3.73)
TU	.084 (3.03)	.084 (3.48)	.068 (2.96)	.133 (4.92)	.091 (2.38)
*TU*ZCOORD*	-.031 (1.68)				
*TU*ZEP*	.020 (2.81)	.019 (3.27)	.021 (4.33)	.026 (2.54)	.065 (1.89)
CBI	-.007 (.12)				
GDP	.027 (1.20)				
OPEN	-.030 (.61)				
*TU*ZCBI*	-.171 (1.35)	-.195 (2.50)	-.187 (2.72)	-.348 (5.03)	-.063 (.40)
*TU*ZGDP*	-.106 (2.56)	-.079 (2.81)	-.068 (2.53)	-.020 (1.43)	-.142 (2.95)
*TU*ZOPEN*	.045 (.55)				
*TU*ZCOORD*ZOPEN*	-.061 (1.13)	-.109 (2.37)	-.140 (3.32)	-.163 (3.10)	-.148 (3.71)
*TU*ZCOORD*ZCBI*	-.007 (.06)				
*TU*ZCOORD*ZGDP*	-.020 (.71)				
OIL				.125 (12.17)	
*OIL*ZCOORD*			-.055 (4.00)	-.060 (3.69)	-.107 (4.23)
*OIL*ZCOORD*ZOPEN*			-.200 (2.95)	-.232 (3.24)	-.147 (2.05)
Fixed effects	Yes	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes
No. of countries	20	20	20	20	20
No. of observations	136	136	136	136	79
Root mean square error	0.01773	0.01713	0.01649	0.02256	0.01856
AR(1) error parameter	0.01	0.04	0.05	-0.03	-0.286

Regressions for 5 year inflation rates in 20 countries, over 1961-95 in models (1)-(9) and 1976-95 in (10). Inflation measured as a decimal, .01 approximately corresponds to 1%. Estimation by feasible generalised least squares allowing for heteroscedastic AR(1) errors. Coefficient estimates reported, absolute t-ratios in brackets.

Table 3: The effects of a 10 percentage point increase in the unionisation rate

VARIABLE	COEFFICIENT	STD DEV OF ZERO MEAN PART	COEFF*STD DEV*0.10
TU	0.068	-	0.0068
*TU*ZEP*	0.021	0.592	0.0013
*TU*ZGDP*	-0.068	0.367	-0.0025
*TU*ZCBI*	-0.187	0.158	-0.003
*TU*ZCOORD*ZOPEN*	-0.14	0.147	-0.0021

Coefficients based on col. (8) in Table2. Scaling factor of 0.10 in final column corresponds to 10 percentage point increase in unionisation. Numbers in final column need to be multiplied by 100 to give impact on percentage inflation rate. Results are approximate because they use the assumption log (1+x)~ x.

Table 4

Robustness tests of the core OECD inflation equation

Regression	1	2	3	4	5
Constant	.017 (1.27)	-.001 (.07)	.009 (1.37)	.004 (.23)	.008 (.88)
INF (-1)			.575 (7.87)	.082 (.83)	.463 (7.77)
TU	.032 (1.60)	.074 (2.05)	.021 (2.00)	.100 (2.62)	.027 (1.89)
*TU*ZEP*	.019 (3.13)	.020 (2.73)	.013 (1.61)	.020 (1.24)	.020 (2.51)
*TU*ZCBI*	-.161 (2.60)	-.150 (1.74)	.001 (.03)	-.158 (1.30)	.041 (.86)
*TU*ZGDP*	-.055 (2.41)	-.081 (2.46)	-.063 (3.06)	-.098 (2.28)	-.109 (3.66)
*TU*ZCOORD*ZOPEN*	-.109 (3.06)	-.135 (2.80)	-.014 (.36)	-.119 (1.99)	-.053 (2.00)
*OIL*ZCOORD*	-.065 (4.68)	-.061 (3.23)	-.092 (4.11)	-.063 (2.72)	-.103 (4.36)
*OIL*ZCOORD*ZOPEN*	-.214 (3.50)	-.251 (3.00)	-.190 (1.86)	-.262 (2.63)	-.291 (4.18)
Estimation method	FGLS	OLS	OLS	OLS	GMM-SYS
Fixed effects	Yes	Yes	No	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes
No. of countries	20	20	20	20	20
Time dimension	1961-98	1961-95	1961-95	1961-95	1961-95
No. of observations	153	136	117	117	117
AR(1) error parameter	0.06	-	-	-	-
Root mean square error	0.0163	0.01629	0.01921	0.01669	0.02025 *
For GMM-SYS estimate					
Sargan test	-	-	-	-	10.26, p=1.00
AR(1) test	-	-	-	-	-3.51, p=.00
AR(2) test	-	-	-	-	-0.99, p=.32

* This calculation is based on residuals from fitting the levels part of the estimator.

Regressions for 5 year inflation rates in 20 countries. Inflation measured as a decimal, .01 approximately corresponds to 1%. Estimation by feasible generalised least squares allowing for heteroscedastic AR(1) errors, OLS and GMM-SYS. GMM-SYS obtained by 1-step estimation, Sargan test computed using 2-step estimation, see Blundell and and Bond (1998). Coefficient estimates reported, absolute t-ratios in brackets, column (2) t-ratios use White (1980) standard errors.

Table 5

Further determinants of inflation in the OECD

Regression	1	2	3
Constant	.002 (.11)	-.022 (1.19)	.009 (.58)
TU	.0708 (2.93)	.092 (2.83)	.047 (2.00)
*TU*ZEP*	.027 (4.19)	.008 (.41)	.024 (4.55)
*TU*ZCBI*	-.204 (3.10)	-.193 (2.32)	-.121 (1.58)
*TU*ZGDP*	-.067 (2.33)	-.026 (.76)	-.057 (2.09)
*TU*ZCOORD*ZOPEN*	-.147 (3.30)	-.158 (3.54)	-.159 (4.05)
*OIL*ZCOORD*	-.057 (4.17)	-.044 (3.24)	-.056 (4.08)
*OIL*ZCOORD*ZOPEN*	-.201 (3.12)	-.166 (2.42)	-.212 (3.24)
LEFT	.005 (.35)		
RIGHT	-.007 (.51)		
*TU*ZLEFT*	.009 (.34)		
*TU*ZRIGHT*	.022 (.65)		
*TU*ZFOI*		-.113 (3.54)	
*TU*ZHOME*			.128 (1.58)
Fixed effects	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes
No. of countries	18	16	19
No. of observations	126	112	133
Root mean square error	0.01555	0.01627	0.01619
AR(1) error parameter	0.05	0.05	0.08

Regressions for 5 year inflation rates in 20 countries. Inflation measured as a decimal, .01 approximately corresponds to 1%. Feasible generalised least squares (FGLS) allows for heteroscedastic AR(1) errors. Coefficient estimates reported, absolute t-ratios in brackets.