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# TRADE UNION DENSITY AND INFLATION PERFORMANCE: EVIDENCE FROM OECD PANEL DATA

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# Trade union density and inflation performance: evidence from OECD panel data

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**Abstract**

This paper examines the impact of union membership rates on inflation in OECD countries. A positive effect of union density is estimated, even after controlling for fixed effects and time dummies. Additional institutional characteristics, for example union coordination, employment protection laws and central bank independence, do not affect inflation directly in a panel setting, but do influence the size of the unionisation coefficient via interaction terms. The results are robust to controlling for potential common causes such as oil price shocks and the political stance of the government, and to using GMM/IV techniques to handle possible endogeneity biases.

*Keywords:* Inflation, trade union density, labour markets, central bank independence, OECD economies.

*JEL classification:* E31, J51.

# I INTRODUCTION

Explaining inflation outcomes is an important research topic in macroeconomics. Theoretical models such as that proposed by Barro and Gordon (1983) predict that equilibrium inflation depends on the factors determining a policy-maker's aversion to inflation, for example the conservativeness of the central bank (Rogoff (1985)) and openness to trade (Romer (1993), Lane (1997)). Recent contributions to this literature emphasise the role of labour markets, for example Cukierman and Lippi (1999) present a model in which the extent of trade union centralisation interacts with central bank independence in setting equilibrium inflation.

Empirical evidence on the determinants of inflation has been presented in a number of well known contributions. Cukierman (1992) and Alesina and Summers (1993) identify a robust link between measures of central bank independence (CBI) and inflation. Romer (1993) estimates a negative correlation between trade openness and inflation and Cukierman and Lippi (1999) provide preliminary evidence for a hump shaped relationship between inflation and union centralisation. The link between inflation and each of these factors has generally been demonstrated using cross-country regressions or pooled time series regressions that do not control for country fixed effects. As most of the variables emphasised evolve very slowly through time it is not clear that they account for the important shifts in inflation regimes that have occurred during recent decades, e.g. average inflation in the OECD rose from 3.7% during $1961-65$ to 10.6% during $1976-80$ before falling to 3.2% during $1991-95$.

A common view of the time variation in inflation is that it derives from supply shocks such as the oil price hikes and industrial disputes of the 1970s. Although clearly important in accounting for changing inflation performance this interpretation leaves much unexplained, for example why did some countries prove more susceptible to inflationary supply shocks than others? This paper emphasises the role of trade unions in explaining inflation. We report panel data regressions demonstrating a positive effect of unionisation rates on average inflation even after controlling for country fixed effects and time dummies. The size of this effect decreases the more highly coordinated the union sector but increases the more strict is employment protection legislation. We also show that factors that correlate negatively with inflation in the cross-section, for instance per capita income and CBI, are insignificant after controlling for fixed effects and time dummies but do affect the impact of unionisation rates on inflation.

In our empirical analysis we try to move beyond existing work in several ways. We show that our core results are observed in various sub-samples obtained by changing the cross-sectional and time series dimensions of the panel. We then address the possibility that the correlation between union density rates and inflation is incidental and arises only as a by-product of other macroeconomic events, e.g. oil price hikes may raise inflation and prompt workers to join unions in search of job security, or governments from the political right may choose to simultaneously reduce union membership and inflation. We provide evidence against these hypotheses through demonstrating that our results are robust to controlling for oil price shocks and a measure of the political stance of the government. Generalised method of moments (GMM) and instrumental variables (IV) estimates of our preferred regression specification are also presented in order to address potential reverse causation biases. The results confirm the positive effect of union density on inflation although some of the interaction terms are less robust.

Our paper is related to work by Blanchard and Wolfers (2000) and Nickell et al (2005), who examine the effect of labour market institutions on unemployment, and Nunziata (2005) who studies the link between institutions and real wages. These contributions suggest possible channels linking union density and inflation. If high unionisation rates are associated with monopoly power in labour supply real wages may be set above competitive levels and the equilibrium unemployment rate will be high. This in turn could induce policy-makers to launch surprise inflation episodes more frequently in order to deliver temporary reductions in unemployment, as in Barro and Gordon (1983), but ultimately this serves only to raise average inflation. Alternatively, labour market structures may influence average inflation independently of the natural rate of unemployment and equilibrium real wages. Ball (1995) presents a model in which the real equilibrium is unique and the inflation preferences of policy-makers are unknown. In the pooling equilibrium both weak and tough policy-makers pursue low inflation in steady-state, but a positive shock to inflation can induce the separating equilibrium because only the tough policy-maker chooses to return inflation to target. If strong trade unions are a source of inflation shocks, e.g. because they periodically attempt to secure higher wages, the preferences of weak policy-makers will be revealed more often and equilibrium inflation will be higher.

The remainder of the paper expands on these points and is structured as follows. Section 2 briefly 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.

# II MODELS OF INFLATION PERFORMANCE

Barro and Gordon (1983) analyse discretionary monetary policy in a model in which equilibrium unemployment exceeds the welfare maximising level. In this framework there exists an incentive for the monetary authority to launch inflation surprises, which ultimately raise equilibrium inflation through their effect on private sector expectations. The determinants of inflation are the factors that influence the incentive to create surprise inflation. Rogoff (1985) shows that if monetary policy is set by an independent central bank and if the central bank is more inflation averse than society in general, equilibrium inflation will be lower than if policy had been set by a government acting to maximise social welfare. In order to test this hypothesis Cukierman (1992) and Alesina and Summers (1993) develop indices of the legal independence of central banks. If central bankers are more inflation averse than elected governments it follows that greater CBI reduces average inflation and cross-country regressions are shown to support this prediction.

Open economy extensions of the Barro-Gordon framework have been developed by Romer (1993) and Lane (1997). In both models the net marginal benefit of monetary expansion is negatively related to openness. In the Romer case this is because terms of trade adjustments restrict the output gain associated with policy expansion and these effects are stronger the more open the economy. In the Lane case the wedge between equilibrium and socially optimal employment is smaller in the tradable sector and this means that temporary increases in output and employment yield smaller benefits at higher levels of openness. A policy authority starting from zero inflation has a smaller incentive to implement surprise inflation the more open the economy and this implies a negative relationship between openness and inflation. Both Romer and Lane present evidence from cross-sectional regressions consistent with this hypothesis.

Cukierman and Lippi (1999), hereafter CL, extend the Barro-Gordon framework to consider imperfect competition in the labour market. The main insight is that trade unions set real wages in excess of the competitive level and therefore increase the wedge between equilibrium and socially optimal employment. This increases the time consistent inflation rate for the standard Barro-Gordon reasons. CL emphasise the degree of centralisation in the union sector (the reciprocal of the number of unions) as a determinant of the final inflation rate. Initially, increasing centralisation implies greater monopoly power and hence a larger real wage premium and higher inflation. However, an increase in average union size also implies that each union is more aware of the inflationary consequences of submitting high wage claims (small unions do not make this link because their impact at the macro level is infinitesimal). To the extent that unions dislike inflation, increasing union size will moderate real wage demands and eventually this effect will dominate that from monopolistic wage-setting, such that the real wage premium and hence inflation will be hump shaped in the degree of centralisation (this is the inflation equivalent of the Calmfors and Driffill (1988) hypothesis concerning unemployment and labour market centralisation). CL show that the turning point of the hump occurs at a higher level of centralisation the greater is CBI. The reason is that high levels of CBI lead unions to believe that inflation will be kept low even if wages are set high and this weakens the second of the two effects.

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 union centralisation, plus interactions between the dummies and the Cukierman index of CBI. The results support the theoretical predictions made by CL.

**The role of unionisation rates**

Another labour market institution that may affect inflation performance is the percentage unionisation of the workforce, often referred to as trade union density. A high level of union density implies reduced substitutability of union and non-union labour and this may shift bargaining power from firms to workers in the labour market. If real wages are set higher as a result, equilibrium unemployment will rise and the time consistency problem facing policy-makers will be more severe. This is a potential channel linking union density and inflation. An alternative mechanism is that some policy-makers accommodate increases in inflation induced by wage hikes associated with strong unions. This implies that otherwise temporary bouts of wage inflation can induce long-lived movements in inflation. As noted in the introduction, Ball (1995) presents a version of the monetary policy game in which central bank preferences are unknown. In the pooling equilibrium both weak and tough policy-makers pursue low inflation in steady-state but a positive shock to inflation can induce the separating equilibrium because only the tough policy-maker chooses to return inflation to target. If strong trade unions are a source of frequent wage hikes the preferences of weak policy-makers will be revealed more regularly and equilibrium inflation will be higher.

The empirical evidence linking inflation and union density has been considered by Hall and Franzese (1998), Chou (2000) and Daniels, Nourzad and VanHoose (2003). A positive effect of union density on inflation is reported in each case but important aspects of the relationship are not examined. Firstly, although the analyses are based on panel data, fixed effects and time dummies are not included in all of the regressions and therefore it is not clear that trade union density is a robust determinant of inflation performance through time. Secondly, potential interactions between union density and other labour market institutions are not explored. One possible interaction is that unionisation rates exert a smaller effect on inflation if the degree of union coordination is high (coordination refers to the extent to which unions communicate during the bargaining process and hence are able to take account of the macroeconomic consequences of their decisions). This could be the case if highly coordinated unions believe that excessive real wages lead to high unemployment and inflation, and choose to exercise restraint in order to avoid these outcomes. Another possibility is that relatively strict employment protection laws increase the positive effect of union density on inflation through transferring bargaining power to the side of the union. Thirdly, previous work has not investigated the hypothesis that the relationship between union density and inflation is the result of a common cause, or reverse causation biases. In this paper we try to cast some light on each of these issues.

## III A PANEL MODEL FOR INFLATION IN THE OECD

The data that we consider cover 20 countries observed over the period $1961-95$ (some series are unavailable post-1995 but for those variables that are available for a longer period we report models for an extended sample in our robustness section).[1] 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).

• *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 shows greater time variation than the original Cukierman index.

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

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

• *TU* is the proportion of employees that belong to a trade union. The feasible range for this variable is $0 - 1$.

• *COORD* measures 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 amongst labour unions. This variable is closely related to the measure of centralization in union bargaining used by CL, since centralisation is the main determinant of the degree of coordination that can be achieved. The *COORD* index can be thought of as a *de facto* measure of the degree of coordination while centralisation is a *de jure* measure. The former provides a better measure of the extent to which unions are likely to moderate wage demands in order to stabilise macroeconomic performance, because it refers to actual as opposed to potential coordination.

• *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 Figure 1 we plot the cross-country averages of *INFLATION*, *TU*, *COORD*, *EP*, *CBI*, *OPEN* and *GDP* at the annual frequency. The units are those described above (note that in Figure 1 the log transform has not been applied to the *GDP* series). 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.

**Insert Figure 1 about here.**

**Econometric methodology**

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

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

In this notation $i$ refers to a country and $t$ to a series of non-overlapping 5 year periods starting $1961-65$ and ending $1991-95$. The dependent variable, $INF$, is defined as $\ln(1 + INFLATION)$. This transformation downweights the importance of any high inflation outliers without exaggerating the effect of low inflation observations. The vector $\mathbf{x}_1$ comprises the determinants of inflation aversion, $\mathbf{x}_2$ is a vector of labour market institutions indicators, $\mathbf{h}$ is a vector of interactions between the first two sets of variables, $\mu_i$ a country fixed effect, $\lambda_t$ a time dummy that controls for factors affecting world inflation and $\varepsilon_{it}$ the error term.

Inflation is measured as an average over a 5 year period in order to capture the equilibrium inflation rate within a macroeconomic regime. This quantity is preferred to the annual inflation rate because the hypotheses that we test relate to equilibrium inflation outcomes rather than high frequency inflation fluctuations that will depend on the cyclical and supply-side conditions that apply at a particular point in time. The phase-averaging approach that we use here has often been employed in testing positive theories in macroeconomics, e.g. in the case of inflation (Gruben and McLeod (2004)), unemployment (Blanchard and Wolfers (2000)) and growth (Levine et al (2000)). However, we recognise that the method has been criticised in some parts of the literature, see for example Hendry and Ericsson (1991), and therefore in the robustness section we show that results estimated using data at the annual frequency are qualitatively similar to those obtained by estimating equation (1).

It will be noted that equation (1) does not control for variables that are often included in time series studies of inflation such as unemployment and import prices. The reasons for this are twofold. Firstly, the fact that we concentrate on the determinants of average inflation over the course of a macroeconomic regime means that cyclical variables such as unemployment are less important. Secondly, some key potential determinants of inflation, for example CBI and openness, are thought to affect inflation through setting the incentive to launch surprise inflation and the strength of this incentive will influence reduced form macroeconomic variables such as unemployment and import price inflation. In order to estimate the full effect of monetary policy and labour market institutions on inflation we do not hold constant these intermediate variables in our main results. However, in our robustness section we control for a measure of detrended unemployment and show that our main findings are unchanged.[2]

In order to test the hypothesis that a variable $x_{m2}$ interacts with a variable $x_{n2}$ in setting inflation, we use the terms $\gamma_{m2} \cdot x_{m2} + \gamma_{m3} x_{m2} \cdot x_{n2}$ in the regression, and define $x_{n2}$ so that it has a zero mean across the sample. This ensures that the coefficient $\gamma_{m2}$ can be interpreted as the coefficient of the "average" country, i.e. the country characterized by the sample average value of $x_{n2}$. In the results section a variable preceded by $Z$ is in zero mean form.

A final point to note is that equation (1) 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). Results obtained using alternative assumptions concerning the error distribution are reported later in the paper.

# IV 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. The objective of this exercise is to demonstrate that our data yield relationships between inflation, central bank independence, openness and income per capita broadly similar to those reported in previous studies that have concentrated on cross-country relationships. In column (1) of Table 1 the explanatory variables are *CBI*, *OPEN* and *GDP*. Each of the variables is negatively signed and significant, in line with the results from past research.

**Insert Table 1 about here.**

In column (2) we consider the CL hypothesis that inflation is hump shaped in union centralisation and that the location of the hump depends on *CBI*. As noted in section 3, *COORD* is closely related to the centralisation index used by CL and can be used in testing the hump shape hypothesis. In column (2) the explanatory variables are *ZCOORD*, *ZCOORD*ZCOORD* and the interactions between those two variables and *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 and the evidence is somewhat weaker on controlling for *CBI*, *OPEN* and *GDP* in column (3).[3]

In column (4) we add fixed effects to the column (1) specification. *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 shifts in the inflation regime 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.

**A detailed look at 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 union density. The coefficient on the unionisation rate is positively signed and significant at the 1% level, supporting the view that monopoly power in the labour market increases equilibrium inflation. In column (2) $EP$ and $COORD$ are added to the regression but neither term is significant at the 5% level. In column (3) $TU$ is interacted with $ZCOORD$ and $ZEP$. It appears that high levels of coordination moderate the inflation increasing effect of highly unionised labour markets whilst employment protection above average makes the effect stronger. In column (4) we show that conditioning on union density does not restore the significance of $CBI$ and $OPEN$ within a panel setting.[4]

**Insert Table 2 about here.**

In column (5) trade union density is interacted with zero mean versions of $CBI$, $GDP$ and $OPEN$, and in column (6) a tested down version of equation (5) is reported. This is obtained by 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 union density. The interaction terms indicate that a relatively high level of coordination reduces the impact of union density on inflation, possibly because large unions recognise that excessive wage claims translate into high inflation and choose to moderate pay demands in order to avoid this outcome. In contrast, the level of union centralisation/coordination does not affect average inflation directly. Furthermore, adding the variables *ZCOORD, ZCOORD*ZCBI, ZCOORD*ZCOORD* and *ZCOORD*ZCOORD*ZCBI* to column (6) leaves the significance of the existing regressors unchanged but does not provide any support for the CL hypothesis (full results are available on request). This finding is consistent with that of Hall and Franzese (1998), and we note that the theoretical underpinnings of the CL hypothesis have been questioned by Fracasso and Ozkan (2004) and Ciccarone and Marchetti (2002).

The effect of union density on inflation appears to increase with the strictness of employment protection legislation, although the statistical significance of this result is less strong in column (3) than in column (6). One interpretation of this result is that high levels of employment protection transfer bargaining power to unions and enable them to extract higher real wages. As argued in our introduction, these wage increases may raise inflation directly or push up equilibrium unemployment and create an incentive for monetary authorities to adopt more expansionary policies that subsequently increase inflation.

Central bank independence above the OECD average reduces the impact of union density on inflation. One interpretation of this result is that unions believe that independent central banks are committed to maintaining stable inflation. Therefore the threat that a policy-maker will respond to high wage demands by tightening policy and driving unemployment to a level that unions cannot tolerate is more credible, and as result wage-setting is less aggressive. The German experience provides a possible example of this interaction. During the period in which the Bundesbank controlled monetary policy in Germany its credible anti-inflation stance appears to have moderated the propensity for trade unions to demand wage increases in excess of productivity growth, and this union restraint contributed to low inflation conditions in Germany, see Soskice and Iversen (1998).

An above average level of income per capita reduces the impact of union density on inflation. This implies that the effect of higher union density on inflation is smaller in countries at the top end of the income distribution, e.g. the United States and Switzerland, than in countries at the bottom end, e.g. Portugal. Several interpretations of this finding are possible. It could be the case that unions are less inclined to push for increases in wages if living standards are already relatively high. A related idea is that unions are more militant in countries in which an independent union movement does not have a long history and such countries tend to be those at the lower end of the OECD income distribution, e.g. in Portugal and Spain per capita income is below the cross-sectional sample average and in those countries trade unions were not free of legal constraints until the transition to democracy in the 1970s.

The fact that incomes have trended up over time implies that union density exerted a stronger effect on inflation performance at the start of the sample than at the end. To be precise, the average value of $ZGDP$ during $1961-65$ was $-.456$ which implies that the coefficient on $TU$ was .150 for a country in the middle of the income distribution, whilst during the period $1991-95$ the average value of $ZGDP$ was .332 which implies that the coefficient on $TU$ was .051 for the country in the middle of the income distribution. It is possible that the powerful effect associated with this interaction term partly reflects a tendency for labour market conditions to exert a weaker effect on average inflation during the post-1980 period as a result of policy-makers focussing on the objective of low inflation and adopting less accommodating policy in response to real wage hikes. However, if the union density term is interacted with a post-1980 dummy the resulting interaction is insignificant, while each of the other variables in the model, including $TU * ZGDP$, remains significant. The role of higher income levels in reducing the effect of union density on inflation through time is reflected in the fact that if $TU_{it} * ZGDP_{it}$ is replaced by $TU_{it} * Z(GDP_{it}/\overline{GDP_t})$ where $\overline{GDP_t}$ is average GDP across countries in period $t$, the $TU$ coefficient falls to .046 (absolute *t-ratio* is 1.94), i.e. removing the drift from the $ZGDP_{it}$ term reduces the importance of the $TU$ effect.[5]

One explanation for the $TU$ effect getting weaker as GDP per capita rises is that unions' wage aspirations do not keep pace with productivity improvements, which means that the inflation pressures associated with wage push factors ease through time. This idea has been used by Ball and Moffitt (2002) in explaining US inflation, although the lag between productivity and wage aspirations is temporary in their analysis.

It is interesting to note that Chou (2000) and Daniels, Nourzad and VanHoose (2003) estimate a hump shaped relation between union density and inflation, which suggests that as unionisation rates increase the marginal effect of inflation declines and conceivably turns negative. If the square of union density is added to model (6) in Table 2 it is negatively signed but marginally insignificant at the 5% level. The estimated coefficients of the other regressors are robust, however, with the one exception of $TU * ZCBI$ for which the coefficient is $-.130$ (*absolute* $t = 1.73$). It could be the case that it is not high unionisation rates *per se* that moderate the marginal effect of union density on inflation but rather that high unionisation rates tend to occur alongside factors that limit the inflation increasing effect of union density, e.g. above average levels of labour market coordination.

**The quantitative significance of the results**

A hypothetical country that is exactly at the sample average in terms of $EP$, $GDP$, $CBI$ and $COORD$ faces an increase in equilibrium inflation of 0.93 percentage points following a 10 percentage point increase in union density (using the column (6) results). This effect is very small, but it is important to bear two points in mind. Firstly, the model controls for common global trends in inflation and union density and therefore provides 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 (6). This yields a coefficient on union density of 0.24, implying that a 10 percentage point increase in unionisation raises equilibrium inflation by approximately 2.4 percentage points in the 'average' country. Secondly, changes in inflation could be much larger given a particular institutional configuration, e.g. high levels of employment protection and low levels of coordination.

A related question is whether or not the marginal effect of union density is always positive given that many interaction terms enter the model. In Table 3 we list the changes in average inflation associated with each regressor following a 10 percentage point increase in unionisation (these calculations are based on column (6) in Table 2). It is clear that after controlling for differences in the variability of the interaction terms it is very unlikely that the derivative of inflation with respect to union density turns negative - setting *ZCOORD*, *ZGDP* and *ZCBI* one standard deviation above zero and *ZEP* one standard deviation below zero gives a positively signed unionisation effect.

**Insert Table 3 about here.**

**Robustness and sensitivity**

In Table 4 we consider the stability of the preferred inflation equation. The first possibility that we investigate is that the impact of institutional factors on inflation has changed over time, e.g. as a result of the breakdown of the Bretton Woods fixed exchange rate system. This involves estimating the specification from Table 2, column 6 for the sub-sample of observations beginning in $1976-1980$. The results (Table 4, column 1) show that most of the relationships are robust. The $CBI$ interaction loses significance, suggesting that in the full sample the effect of $CBI$ is mainly due to events in the 1960s and early 1970s. An inspection of the data reveals 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 to support new exchange rate targets (consider the experience of the United Kingdom in 1967 when the Wilson government chose to devalue sterling). These reductions in $CBI$ appear to have amplified the inflation increasing effect of rising union density during the $1960s$ and $1970s$ and therefore explain a large part of the negative point estimate on *TU*ZCBI* in the full sample regression.

**Insert Table 4 about here.**

The second column of Table 4 provides evidence on the robustness of the basic conclusions when the time dimension of the panel is extended. This involves adding an extra set of observations based on three year averages for $1996-98$ (most of the series end in 1998). The final observations for union density for Canada and Spain are $1996-97$ averages, whilst the final observation for Belgium is missing. The *CBI* variable is unavailable post-1995 and therefore 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 (omitting the term in CBI yields similar results to those that we report). The coefficient on union density remains significant but falls by roughly one third, suggesting that recent inflation performance has been less highly correlated with movements in union density (though bear in mind the limitations of the data).

The cross-sectional stability of the results is evaluated in Figure 2, which plots coefficient estimates for the preferred model (Table 2, column 6) obtained by deleting one country at a time from the panel. The stability of the coefficients is confirmed in most cases, though there are some exceptions. The interaction between union density and coordination falls by one half when Finland is excluded from the sample, whilst the coefficient on the employment protection interaction loses significance when Sweden is excluded, mainly because it is imprecisely estimated. These are isolated episodes, however, and overall the plots do not suggest that our core results are due to outlying observations associated with a particular country, or the pooling biases that would arise in the event of coefficient heterogeneity across countries.

**Insert Figure 2 about here.**

Columns (3) and (4) in Table 4 report estimates of the preferred specification obtained by OLS rather than FGLS. The column (3) results are based on standard errors calculated using the Huber-White method, whilst the column (4) results are based on panel corrected standard errors that adjust for cross-sectional correlations in the residuals that vary across country pairs (common factors in the residual series are handled by the time dummies). The significance of the coefficient estimates is generally robust, though the method based on panel corrected standard errors points to some uncertainty in the estimation of the coefficient for the interaction between union density and coordination.

The final two columns of Table 4 contain estimates based on annual data rather than the method of phase averaging. Column (5) reports a static model and column (6) reports a dynamic model that accounts for inflation persistence by including once lagged inflation. The latter specification also includes the deviation of unemployment from a Hodrick-Prescott trend in order to control for cyclical movements in inflation (the Hodrick-Prescott smoothing parameter is set to 400, the recommended value for annual data). Both models contain a full set of fixed effects and time dummies and reported t-ratios are based on robust standard errors. The results for the static model are broadly in line with those obtained using 5 year averages, the exception being the $TU * ZCBI$ term. In the dynamic model the coefficients are generally smaller in absolute terms, but after multiplying each of them by $(1/(1 - .578)) = 2.37$ in order to account for the dynamic propagation implied by the autoregressive term the effects are close to those obtained previously. The union density coefficient is significant at only the 15% level, possibly reflecting additional variation in annual inflation that is not fully captured by the unemployment term, and the interaction featuring $CBI$ is insignificant and incorrectly signed (this term proves least robust across the range of models that we estimate). Overall, however, we conclude that our main results are not induced by the method of phase averaging that we have employed.

**Adding further controls**

We now augment our preferred model with additional controls in order to evaluate various interpretations of our results in which the relationship between union density and inflation is spurious or co-incidental. One possibility is that the positive correlation we have estimated arises only because oil price shocks forced inflation up during the 1970s and also prompted workers to join trade unions in search of protection against macroeconomic turbulence. As oil price shocks are typically global events one might expect their effects to be captured by the time dummies, but there is no guarantee of this, e.g. if there are multiple global shocks the time dummies will pick up their average effect and some oil price movements will remain in the residuals. In column (1) of Table 5 we control for the natural log of one plus the annual rate of increase of the US$ spot price of a barrel of West Texas Intermediate (denoted $OIL$). As this variable takes the same values across countries it is collinear with the time dummies and therefore the first time dummy is omitted to facilitate estimation. The measure of oil shocks is highly significant, but this does not affect the role of labour market variables, each of which is remarkably robust.

**Insert Table 5 about here.**

The next hypothesis that we address is that changes in the political persuasion of the government cause both inflation and union density - consider the case of the UK in the 1980s in which the Thatcher administration implemented legislation that reduced union membership and also adopted restrictive fiscal and monetary policies that brought down inflation. In column (2) we control for $LEFT$ and $RIGHT$, which measure the fraction of cabinet seats taken by parties from the political left and political right respectively (the series are not collinear because some seats are taken by centrist parties).[6] The variable $LEFT$ is positively signed, suggesting that left wing governments are associated with higher inflation but the t-ratio for this term is just 1.20, while the coefficient multiplying $RIGHT$ is practically zero. In contrast, the effects based on union density are very robust. In column (3) we allow for interactions between union density and the measures of political stance and obtain similar results for the labour market variables.

Hence, there is little support for the idea that shifts in the political landscape drive our results.[7]

In column (5) we control for the proportion of households that are owner occupiers, $HOME$. Oswald (1996) argues that this variable is inversely related to geographical labour mobility (the costs of moving house are larger for owner occupiers) and can therefore be expected to raise inflation through creating inflexibilities in aggregate supply. Conditioning on this term does not affect our main results, however. Finally, in column (5) we control for each of the additional terms introduced. Once again, the role of union density is very robust.

**Controlling for endogeneity**

In the final part of this section we try to control for the potential endogeneity of labour market institutions. This could arise if positive shocks to inflation prompt workers to join unions in order to lobby for compensating wage increases. Identifying external instruments for the time-varying component of labour market institutions is difficult and therefore we rely on suitably lagged values of union density and other labour market variables in leveraging the exogenous variation in the data. The validity of these instruments can be assessed by means of careful residual diagnostic testing, as we explain below.

The first technique that we employ is the system generalised method of moments (GMM-SYS) estimator due to Arellano and Bover (1995) and Blundell and Bond (1998). In order to see the logic behind this method consider the following model for inflation:

$$INF_{it} = \gamma_0 + \gamma_1 INF_{it-1} + \gamma_2 TU_{it} + \mu_i + \lambda_t + \varepsilon_{it} \tag{2}$$

Additional controls are omitted to avoid clutter in the exposition but are included in our empirical specifications. The autoregressive term in (2) controls for the dynamics that were previously modelled by assuming an AR(1) error structure in the FGLS estimation. Assuming that (a) $\varepsilon_{it}$ is serially uncorrelated and (b) $TU$ is uncorrelated with future values of $\varepsilon$ (endogeneity of $TU$ implies only that it is correlated with current values of $\varepsilon$) consistent estimation of (2) entails differencing to remove the fixed effects and then using lagged levels of $INF$ and $TU$ dated $t-2$ and earlier as instruments in forming a GMM estimator.[8] Further moment conditions can be obtained through assuming a constant correlation between $INF_{it}$ and $\mu_i$ and $TU_{it}$ and $\mu_i$. Arellano and Bover (1995) show that $\Delta INF_{it-1}$ and $\Delta TU_{it-1}$ can be used to instrument the endogenous variables in (2) directly when these conditions hold because first differenced instruments are orthogonal to the composite error $\mu_i + \varepsilon_{it}$.[9] Blundell and Bond (1998) show that these additional moment conditions yield large efficiency gains when the data exhibit high time series persistence, as is the case for the macro variables considered here. A final technical point is that first differencing (2) in order to use the first set of instruments yields an MA(1) error structure even though $\varepsilon_{it}$ is serially uncorrelated. Hence, if the model is well specified we expect to find evidence of negative first order error autocorrelation but no evidence of second order error autocorrelation. The standard errors take account of the MA(1) error structure, see Arellano and Bond (1991).

The results are reported in Table 6. In column (1) the instrument set is $\{INF_{it-1}$, $INF_{it-2}$, $TU_{it-2}$, $TU_{it-3}$, $TU_{it-4}$, $\Delta TU_{it-1}\}$ plus a full set of time dummies, which are included in each set of instruments used in Table 6. The term $\Delta INF_{it-1}$ is excluded from the instruments because its inclusion led to an autoregressive parameter very close to the least squares estimate, a sign that the instrument is not valid, see Blundell and Bond (1998). The $TU$ coefficient is of comparable magnitude to that obtained by FGLS (see Table 2, column (1)) and is significant at the 7% level. The autoregressive parameter is roughly half the size that estimated in Table 4 reflecting the fact that inflation is less persistent at the 5 year frequency than the annual frequency. The error autocorrelation tests are consistent with the hypothesis that $\varepsilon_{it}$ is serially uncorrelated and the Sargan test with instrument validity (these outcomes are discussed in more detail below).

**Insert Table 6 about here.**

Column (2) adds $TU * ZCOORD$ and $TU * ZEP$ to the model and $\{COORD_{it-2}$, $EP_{it-2}$, $\Delta COORD_{it-1}$, $\Delta EP_{it-1}\}$ to the instruments. Union density is significant at the 5% level, as is its interaction with coordination. The interaction between unionisation and employment protection is insignificant, however, matching the FGLS estimate of the equivalent specification in Table 2. In column (3) $TU * ZCBI$ and $TU * ZGDP$ are added to the model and $\{CBI_{it-2}$, $GDP_{it-2}$, $\Delta CBI_{it-1}$, $\Delta GDP_{it-1}\}$ to the instruments. The positive effect of union density is confirmed and is moderated by high levels of coordination and above average per capita income. The interactions featuring employment protection and CBI are insignificant, however. Column (4) omits the insignificant autoregressive term from (3) and drops $INF$ from the instruments.

The parameter estimates are robust to these changes.

In the final column of Table 6 we address a key objection to the GMM-SYS results, which is that the consistency of the estimator relies on the cross-sectional dimension of the panel being 'large'. In a well known application of this technique to macro panel data Levine et al (2000) consider 63 countries whereas we consider only 20 countries, and this may lead to small sample biases. In order to check this possibility we report in column (5) an instrumental variables estimate of the static model in column (4). We maintain our assumptions concerning regressor endogeneity and therefore use the instruments listed below.[10]

$$\{TU_{it-1},\ (TU*ZCOORD)_{it-1},\ (TU*ZEP)_{it-1},\ (TU*ZCBI)_{it-1},$$
$$(TU*ZGDP)_{it-1},\ ZCOORD_{it-1},\ ZEP_{it-1},\ ZCBI_{it-1},\ ZGDP_{it-1}\}$$

The inefficiency of the IV estimator relative to GMM is reflected in the larger coefficient standard errors. Nevertheless the effect of union density is similar to that obtained previously and is significant at the 5% level. The interaction featuring $COORD$ is insignificant in this case, possibly reflecting the difficulties in instrumenting a potentially endogenous regressor using its own lags when the variable in question exhibits relatively little time variation. Otherwise, the picture is similar to that in columns (1)-(4) in that $TU*ZGDP$ is significant while $TU*ZEP$ and $TU*ZCBI$ are insignificant.

Overall, the results from GMM and IV estimation indicate that the positive effect of union density on inflation is unlikely to be the result of reverse causation bias. The findings that union density exerts a smaller effect on inflation the higher is union coordination and per capita income are also generally robust. On the other hand the interaction terms featuring $EP$ and $CBI$ are less robust, reflecting either that these effects were partly endogenous or that the instruments available explain only a small fraction of the exogenous variation in these variables.

**How appropriate are the instruments?**

A sceptic may argue that the significance of union density in Table 6 arises only because macroeconomic shocks in $t-1$ raise $TU$ in the same period but raise inflation with a delay, i.e. in period $t$. Given some persistence in the $TU$ series even the GMM and IV estimates would then be spurious. The Sargan statistics cast doubt on this hypothesis, however. These statistics are based on the sample analogues of the over-identifying restrictions implied by the instrument set. The fact that each of them is insignificant at the 10% level indicates that the instruments do not influence inflation through channels other than the labour market variables that have been included in the model. A possible objection to this evidence is that it reflects a Type II error arising from the low power of the Sargan test when the number of moment conditions is large, see Bowsher (2002). However, we note that the hypothesis of instrument validity cannot be rejected in the case of the IV estimate reported in column (5) in which there are only 4 over-identifying restrictions. Further, we investigated the consequences of reducing the number of over-identifying moment conditions by omitting one variable at at time from the instruments used in column (4). In each case the evidence for instrument validity remained intact.

The final question that we address is whether or not the chosen instruments have explanatory power for the endogenous variables. The main concern is that if the instruments are weak the parameter estimates may be biased away from zero. In order to check this point we regressed $\Delta TU_{it}$ on each of the levels instruments used in column (4) and $TU_{it}$ on each of the first differenced instruments used in column (4). The $F - statistics$ for the joint significance of the instruments are 16.17 ($p = .04$) and 27.61 ($p = .00$) respectively, suggesting that the instruments are able to identify substantial exogenous variation in union density (full details of these regressions are available on request).

## V SUMMARY

This paper has examined the factors that explain inflation performance in OECD countries. An important theme of the paper was the need to explain shifts in inflation within OECD countries in addition to the cross-sectional differences in inflation that have been the focus of many previous studies. We found that after controlling for fixed effects and time dummies factors such as central bank independence, trade openness and per capita income exert weak effects on inflation performance. In contrast, union density was shown to exert a positive effect on inflation in a panel setting. A possible mechanism underpinning this relationship is that high unionisation rates strengthen the bargaining positions of unions and enable them to extract higher real wages. Such a distortion can raise inflation either through pushing up equilibrium unemployment and inducing central banks to pursue inflationary policy more frequently, or through creating cost-push inflation that is subsequently accommodated by the monetary authority because it will not tolerate the loss in output necessary to reduce inflation. We estimated interaction effects that can be related to this interpretation of the link between unionisation and inflation. In particular, factors that may restrict real wage premia, for example union coordination, central bank independence and relatively high per capita income, were found to decrease the effect of union density on inflation. On the other hand, strict employment protection laws amplify the inflationary effect of unionisation rates.

The second half of the paper examined the robustness of these findings in some detail. Variations in both the cross-sectional and time series dimensions of the sample were considered, and in general the main findings remained intact. We then addressed the possibility that our results are co-incidental in the sense that movements in union membership and inflation are both driven by factors such as oil price shocks or the political stance of the government, but found no evidence that this was the case. Finally, in order to control for potential reverse causation biases, we presented GMM and IV estimates of our core specification in which lagged values of the regressors were used to identify exogenous variation in the relationships of interest. The basic effect of union density proved to be robust, as did the interaction terms featuring coordination and per capita income. The interaction featuring central bank independence and employment protection proved less robust, however, possibly because the limited time variation in these variables means that their lagged values do not serve as good instruments.

## Notes

1. 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.

2. In related work Bowdler and Nunziata (2005) report panel models for annual inflation as a function of lagged inflation, unemployment, import prices, the tax wedge and productivity. It is shown that after controlling for this set of reduced form variables labour market institutions do not affect inflation directly but they do matter for the coefficients with which reduced form variables feed into inflation.

3. 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.

4. 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).

5. It is important to note that as time passes the value of GDP in country $i$ minus average GDP across the entire panel will rise for the later observations and fall for the earlier observations. Therefore in order for the total derivative of inflation with respect to union density to remain positive the coefficient multiplying $TU * ZGDP$ must fall as new time observations are added to the panel. The results that we present in Table 4, column 2 are consistent with this observation.

6. Data on $LEFT$ and $RIGHT$ are not available for Portugal and Spain.

7. Sceptics could point out that the insignificance of $LEFT$ and $RIGHT$ suggests that they are poor measures of political stance and therefore provide only a weak robustness check. However, simple bivariate least squares regressions of $INF$ on $LEFT$ and $RIGHT$ respectively yield significant slope coefficients of the expected sign. This suggests that the insignificance of $LEFT$ and $RIGHT$ in Table 5 is due to labour market variables being the key drivers of inflation rather than $LEFT$ and $RIGHT$ being poor measures of political stance.

8. To be precise, the GMM estimator uses the following moment conditions: $E(INF_{i,t-s}\Delta\varepsilon_{it}) = 0$; $E(TU_{i,t-s}\Delta\varepsilon_{it}) = 0$ for $t = 3, 4, ...T$, and $s \geqslant 2$.

9. In this case the following additional moment conditions are available: $E(\Delta INF_{it-s}(\mu_i + \varepsilon_{it})) = 0$ for $s = 1$ and $E(\Delta TU_{it-s}(\mu_i + \varepsilon_{it})) = 0$ for $s = 1$.

10. Levels and interactions of the zero mean terms are used in order to ensure over-identification of the model. If only the lagged interaction terms are used the model is just identified and the parameter estimates are very similar to those reported in Table 6.

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## Appendix: Data Sources

$INFLATION$ Data are 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.

$CBI$ Data are from van Lelyveld (2000). The index is obtained by aggregating indicator variables describing factors such as the conditions under which central banks extend loans to the government, the terms of reference of the central bank governor. van Lelyveld does not provide data for Portugal and therefore we use the data from Cukierman (1992).

$OPEN$ Data are from the Penn World Tables. German data for the pre-unification period are from the *International Financial Statistics* database, and are spliced to the $1990s$ Penn data to obtain a consistent series.

$GDP$ Data are from 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 $-DM exchange rate. This series was then divided by the US price level and then by German population to give real per capita GDP in US$. Finally, this series is spliced to the Penn series for $1990-98$ to give the data that we use for German GDP.

$EP$ Data are from Blanchard and Wolfers (2000).

$TU$ 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.

$UNEMP$ is based on unemployment data taken from Layard et al (1991) and is updated using the OECD *Employment Outlook 2000.* The Portuguese data are from the London School of Economics CEP-OECD database, and the data for Italy are based on the US Bureau of Labor Statistics series, "unemployment rates on US concepts".

$OIL$ Data are from Dow Jones Energy Service (copyright).

$HOME$ Data are taken from Oswald (1996) and interpolated from a 10 year frequency to a 5 year frequency.

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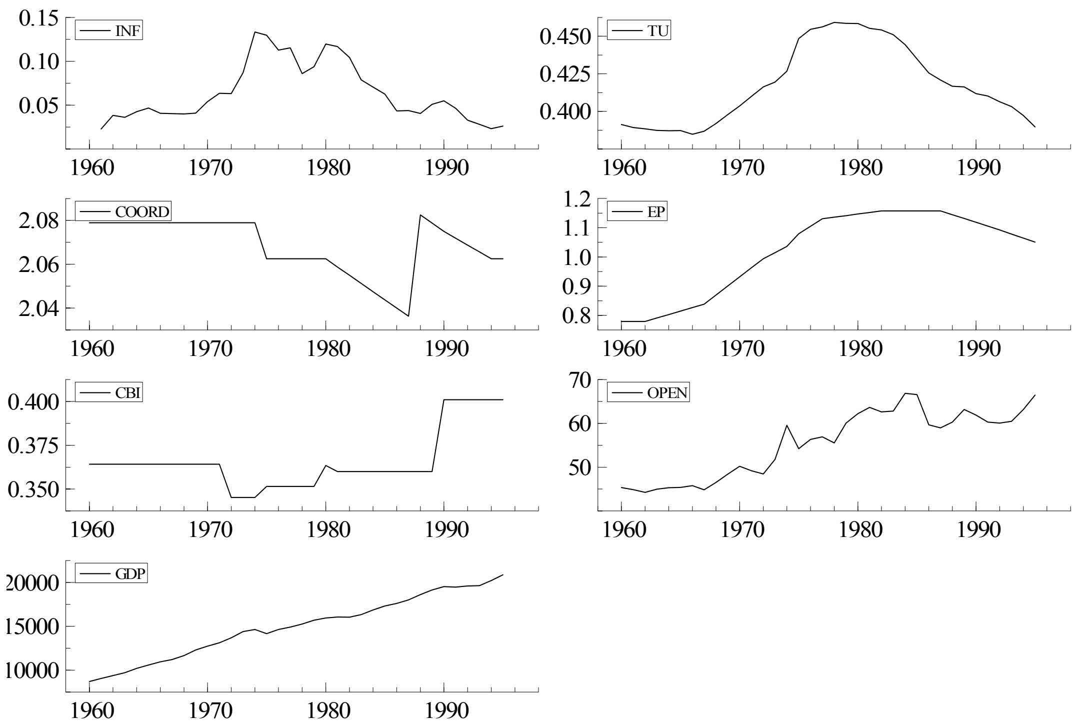


Figure 1: Cross-country averages for *INF*, *TU*, *COORD*, *EP*, *CBI*, *OPEN* and *GDP*.

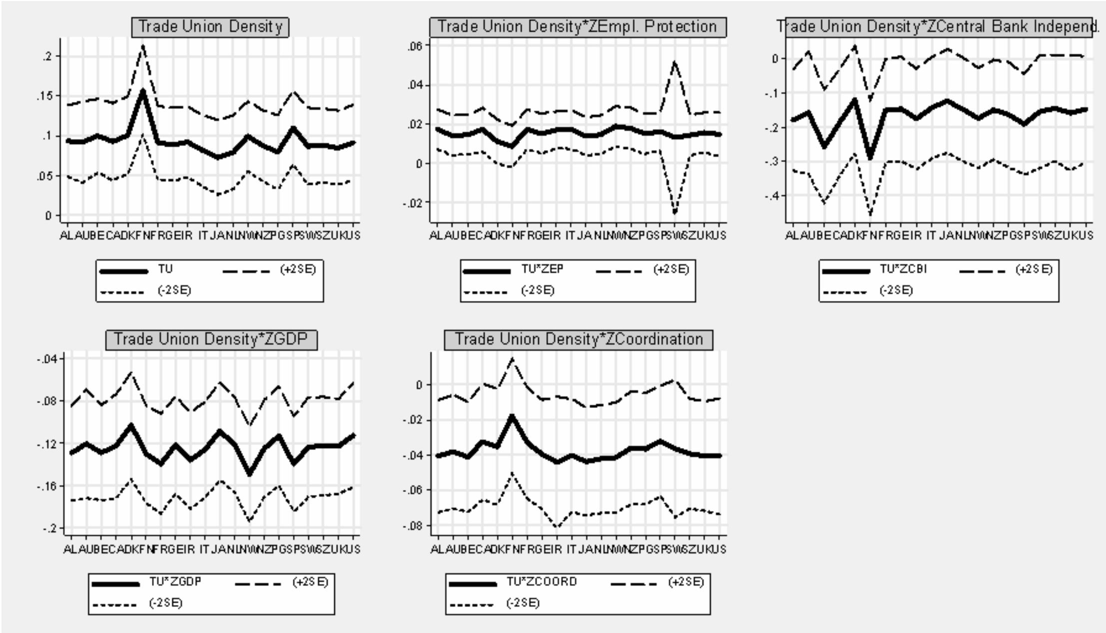


Figure 2: Regression coefficients from table 2, column (6) obtained by dropping one country at a time from the panel.

Table 1

Basic determinants of inflation in the OECD

| Regression | 1 | 2 | 3 | 4 | 5 |
|---|---|---|---|---|---|
| | -.037 | | -.011 | -.025 | -.017 |
| *CBI* | (3.06) | | (0.56) | (0.73) | (0.41) |
| | -.017 | | -.021 | .001 | .011 |
| *OPEN* | (2.42) | | (3.05) | (0.03) | (0.57) |
| | -.038 | | -.038 | -.034 | -.037 |
| *GDP* | (4.34) | | (3.85) | (1.77) | (2.04) |
| | | -.008 | -.009 | | -.006 |
| *ZCOORD* | | (2.28) | (2.84) | | (0.52) |
| | | -.010 | -.013 | | -.023 |
| *ZCOORD*ZCOORD* | | (1.64) | (2.29) | | (1.81) |
| | | .044 | .018 | | .180 |
| *ZCOORD*ZCBI* | | (1.76) | (0.70) | | (2.14) |
| | | -.074 | -.029 | | .067 |
| *ZCOORD*ZCOORD*ZCBI* | | (2.20) | (0.79) | | (0.52) |
| Fixed effects | No | No | No | Yes | Yes |
| Time dummies | Yes | Yes | Yes | Yes | Yes |
| No. of observations | 139 | 136 | 136 | 139 | 136 |
| Root mean square error | 0.0242 | 0.02617 | 0.0223 | 0.0216 | 0.0183 |
| AR(1) error parameter | 0.26 | 0.34 | 0.22 | 0.11 | -0.02 |

Regressions for non-overlapping 5 year inflation rates in 20 countries, 1961-95. FGLS estimation allows for heteroscedastic AR(1) errors. Coefficient estimates reported, absolute t-ratios in parentheses. Regression intercepts are not reported.

Table 2

Labour market institutions and inflation performance in the OECD

| Regression | 1 | 2 | 3 | 4 | 5 | 6 |
|---|---|---|---|---|---|---|
| *TU* | .074<br>(3.25) | .062<br>(2.62) | .048<br>(2.06) | .043<br>(1.89) | .104<br>(4.29) | .093<br>(4.04) |
| *EP* | | .003<br>(0.70) | | | | |
| *COORD* | | -.015<br>(1.52) | | | | |
| *TU*ZCOORD* | | | -.047<br>(2.65) | -.030<br>(1.57) | -.041<br>(2.39) | -.038<br>(2.43) |
| *TU*ZEP* | | | .007<br>(0.98) | .005<br>(0.68) | .016<br>(2.49) | .015<br>(3.07) |
| *CBI* | | | | -.027<br>(0.79) | -.029<br>(0.53) | |
| *GDP* | | | | -.034<br>(1.78) | .025<br>(1.18) | |
| *OPEN* | | | | -006<br>(0.29) | -.041<br>(0.98) | |
| *TU*ZCBI* | | | | | -.126<br>(1.06) | -.169<br>(2.25) |
| *TU*ZGDP* | | | | | -.154<br>(5.16) | -.126<br>(5.57) |
| *TU*ZOPEN* | | | | | .068<br>(0.88) | |
| Fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Time dummies | Yes | Yes | Yes | Yes | Yes | Yes |
| No. of observations | 137 | 137 | 137 | 136 | 136 | 136 |
| Root mean square error | 0.0281 | 0.01883 | 0.01853 | 0.01811 | 0.01758 | 0.0173 |
| AR(1) error parameter | 0.17 | 0.15 | 0.14 | 0.06 | -0.01 | -0.01 |

Regressions for non-overlapping 5 year inflation rates in 20 countries, 1961-95. FGLS estimation allows for heteroscedastic AR(1) errors. Coefficient estimates reported, absolute t-ratios in parentheses. Regression intercepts are not reported.

Table 3

The response of inflation to a 10 percentage point increase in unionisation

| VARIABLE | COEFFICIENT | STD DEV OF ZERO MEAN PART | CHANGE IN INFLATION IF ONE STD DEV ABOVE MEAN |
|---|---|---|---|
| *TU* | 0.093 | - | 0.93 |
| *TU*ZEP* | 0.015 | 0.592 | 0.09 |
| *TU*ZGDP* | -0.126 | 0.367 | -0.46 |
| *TU*ZCBI* | -0.169 | 0.158 | -0.27 |
| *TU*ZCOORD* | -0.038 | 0.147 | -0.06 |

Coefficients based on col. (6) in Table 2. Final column gives the change in the average percentage inflation rate following a 10 percentage point increase in unionisation for a country that is one standard deviation above the sample average for each of the zero mean variables used in the interaction terms.

Table 4

Robustness tests of the OECD inflation equation

| Regression | 1 | 2 | 3 | 4 | 5 | 6 |
|---|---|---|---|---|---|---|
| *TU* | .090<br>(2.84) | .059<br>(2.88) | .095<br>(2.70) | .095<br>(2.99) | .070<br>(2.11) | .033<br>(1.53) |
| *TU*ZCOORD* | -.066<br>(4.19) | -.023<br>(1.62) | -.035<br>(1.97) | -.035<br>(1.21) | -.107<br>(1.92) | -.057<br>(1.92) |
| *TU*ZEP* | .085<br>(2.73) | .011<br>(1.65) | .013<br>(1.79) | .013<br>(2.79) | .066<br>(2.92) | .033<br>(2.35) |
| *TU*ZCBI* | -.059<br>(0.37) | -.174<br>(2.43) | -.167<br>(1.76) | -.167<br>(2.52) | .145<br>(0.64) | .123<br>(1.21) |
| *TU*ZGDP* | -.104<br>(3.23) | -.091<br>(4.73) | -.113<br>(4.07) | -.113<br>(3.12) | -.117<br>(3.15) | -.065<br>(1.98) |
| *INF(-1)* | | | | | | .578<br>(12.5) |
| *UNEMP* | | | | | | -.006<br>(4.50) |
| Estimation method | FGLS | FGLS | OLS-HW | OLS-PCSE | OLS | OLS |
| Fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Time dummies | Yes | Yes | Yes | Yes | Yes | Yes |
| Time dimension | 1976-1995 | 1961-98 | 1961-95 | 1961-95 | 1961-95 | 1961-95 |
| No. of observations | 79 | 154 | 136 | 136 | 686 | 666 |
| AR(1) error parameter | -0.26 | 0.06 | - | - | - | - |
| Root mean square error | 0.01948 | 0.01698 | 0.01702 | 0.01702 | 0.02528 | 0.01925 |

Regressions for non-overlapping 5 year inflation rates in 20 countries (the final set of time observations in column 2 are calculated for 1996-98). Column 6 estimates are based on annual data. FGLS estimation allows for heteroscedastic AR(1) errors. OLS-HW uses Huber-White standard errors, OLS-PCSE uses panel corrected standard errors. Coefficient estimates reported, absolute t-ratios in parentheses. Regression intercepts are not reported.

Table 5

Adding further controls to the OECD inflation equation

| Regression | 1 | 2 | 3 | 4 | 5 |
|---|---|---|---|---|---|
| *TU* | .093<br>(4.04) | .101<br>(4.14) | .093<br>(3.73) | .078<br>(3.33) | .090<br>(3.63) |
| *TU*COORD* | -.038<br>(2.43) | -.033<br>(2.16) | -.034<br>(2.11) | -.038<br>(2.38) | -.036<br>(2.24) |
| *TU*ZEP* | .015<br>(3.07) | .019<br>(3.57) | .021<br>(3.19) | .015<br>(2.89) | .021<br>(3.12) |
| *TU*ZCBI* | -.169<br>(2.25) | -.180<br>(2.42) | -.172<br>(2.28) | -.149<br>(1.90) | -.150<br>(1.93) |
| *TU*ZGDP* | -.126<br>(5.57) | -.127<br>(5.37) | -.126<br>(5.18) | -.112<br>(4.69) | -.127<br>(5.09) |
| *OIL* | .593<br>(4.71) | | | | .617<br>(4.71) |
| *LEFT* | | .007<br>(1.20) | -.004<br>(0.24) | | -.004<br>(0.27) |
| *RIGHT* | | .002<br>(0.30) | -.012<br>(0.92) | | -.011<br>(0.83) |
| *TU*ZLEFT* | | | .028<br>(1.00) | | .029<br>(1.01) |
| *TU*ZRIGHT* | | | .042<br>(1.23) | | .041<br>(1.15) |
| *HOME* | | | | .0001<br>(0.38) | .0002<br>(0.53) |
| Fixed effects | Yes | Yes | Yes | Yes | Yes |
| Time dummies | Yes | Yes | Yes | Yes | Yes |
| No. of countries | 20 | 18 | 18 | 19 | 18 |
| No. of observations | 136 | 126 | 133 | 133 | 126 |
| Root mean square error | 0.0173 | 0.0165 | 0.0165 | 0.0173 | 0.0166 |
| AR(1) error parameter | -0.01 | -0.03 | 0.004 | 0.02 | 0.01 |

Regressions for 5 year inflation rates. FGLS estimation allows for heteroscedastic AR(1) errors. Coefficient estimates reported, absolute t-ratios in parentheses. Regression intercepts are not reported. Models including oil price inflation omit one time dummy to avoid collinearity.

Table 6

GMM-SYS and IV estimates of the OECD inflation equation

| Estimation method | GMM | GMM | GMM | GMM | IV |
|---|---|---|---|---|---|
| *INF (-1)* | .238<br>(2.65) | .243<br>(2.84) | .113<br>(1.16) | | |
| *TU* | .053<br>(1.85) | .054<br>(2.42) | .076<br>(3.18) | .071<br>(2.63) | .110<br>(2.08) |
| *TU*COORD* | | -.055<br>(3.68) | -.033<br>(2.54) | -.031<br>(2.36) | -.008<br>(0.23) |
| *TU*ZEP* | | .008<br>(0.32) | .002<br>(0.12) | .002<br>(0.08) | .031<br>(1.50) |
| *TU*ZCBI* | | | -.051<br>(0.58) | -.081<br>(0.89) | -.063<br>(0.32) |
| *TU*ZGDP* | | | -.136<br>(3.66) | -.123<br>(2.26) | -.139<br>(2.76) |
| Fixed effects | Yes | Yes | Yes | Yes | Yes |
| Time dummies | Yes | Yes | Yes | Yes | Yes |
| No. of observations | 118 | 118 | 117 | 136 | 116 |
| Sargan test statistic | 13.47 | 13.71 | 9.4 | 8.32 | 4.01 |
| Sargan critical value (10%) | 33.2 | 50.66 | 69.92 | 60.91 | 7.78 |
| AR(1) test p-value | 0.009 | 0.002 | 0.011 | 0.002 | - |
| AR(2) test, p-value | 0.471 | 0.486 | 0.285 | 0.246 | - |

Regressions for 5 year inflation rates in 20 countries. Coefficient estimates reported, absolute t-ratios in parentheses. Regression intercepts are not reported. Instruments used are described in the text.