UNEMPLOYMENT BENEFITS AS A SUBSTITUTE
FOR A CONSERVATIVE CENTRAL BANKER

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Abstract—In the many years since their introduction, positive theories of inflation have rarely been tested. This paper documents a negative relationship between inflation and the welfare state (proxied by the parameters of the unemployment benefit program) that is to be expected in such theories. Because unemployment benefits make the monetary authority less concerned about the plight of the unemployed, building a welfare state has a similar effect to appointing a conservative central banker. The relationship holds in a panel of 20 OECD countries over the period 1961–1992, a region where Romer finds no evidence of commitment problems. It holds controlling for country and time fixed effects, country-specific time trends, other covariates, and using a decadal panel. Interpreted as causal, the estimated effect is economically large: a 1-standard-deviation decrease in benefit duration is predicted to add 1.4 percentage points onto inflation, or 31% of the standard deviation in inflation.

I. Introduction

IN 1977 Kydland and Prescott introduced the problem of dynamic inconsistency and used it to analyze the inflation-unemployment trade-off. Barro and Gordon (1983) and Rogoff (1985) later developed the main ideas into a positive theory of inflation. This was deemed particularly relevant in view of the experience of the United States and other countries where the ability to have policy commitment appeared to be insufficient (see Barro & Gordon, 1983, p. 592). Since then, this approach has become standard in the area and is described in virtually every macroeconomics textbook available. In the last 20 years, however, there has been remarkably little empirical research in this area. Romer (1993) is one exception. He finds a negative relationship between openness and inflation for a large cross section of low- and middle-income countries for the post-Bretton Woods period.[1] This is consistent with Barro and Gordon (1983), as there are fewer benefits to surprise inflation in more open economies, particularly when there is a floating exchange rate regime. He finds no relationship between inflation and openness in the high-income sample and argues that these countries have solved the dynamic inconsis-

tency problem, probably through the development of institutions that allow policy commitments.[2]

In this paper we document a strong, negative relationship between the welfare state and inflation in 20 OECD countries for the 32-year period between 1961 and 1992. The association survives the inclusion of country and time fixed effects, country-specific time trends, other covariates, and a lagged dependent variable. The effect of the welfare state (proxied by the parameters of the unemployment benefit program) on inflation is also economically significant: under a causal interpretation, 31% of the inflation variation is explained by variations in our measure of the welfare state. A simple explanation of these findings is in terms of Barro and Gordon (1983). A more generous welfare state makes spells of unemployment less costly. The no-commitment equilibrium rate of inflation is then lower, as the policy-maker is less tempted to inflate. In other words, the welfare state is a substitute for a conservative central banker. In contrast to Romer (1993), our sample is a panel of OECD countries. This means that we cannot reject the hypothesis that dynamic inconsistency problems are still present in developed countries. The paper also discusses the implications of having unemployment benefits respond to economic conditions [as in the evidence presented in Saint Paul (1996), Rodrik (1998), and Di Tella and MacCulloch (1995, 2002)].

In section II we present a simple model and the empirical strategy, and introduce the data we use. In section III we present our results, and section IV concludes.

II. Theory, Empirical Strategy, and Data

A. Basic Theory

Assume a population of identical workers who dislike inflation (π) and live in an economy where there is equilibrium unemployment. As a justification for these assumptions note that inflation is costly because efforts must be made to reduce holding money (which is costless to produce) and that unemployment arises naturally (to help discipline workers) in an economy where there is imperfect monitoring of effort at work. The fraction that are unemployed is denoted u, and they receive benefits b. If employed, they earn a gross wage, W. An individual's (ex ante) welfare is $V = (1 - u) \log (W - T) + u \log (b - T) -$

[1] Terra (1998) has reexamined Romer's findings and claims that they are driven by a group of highly indebted countries. She argues that after the debt crisis, these countries had to generate both the foreign exchange and the public resources to repay them, leading to the observed correlation. Romer (1998), however, shows that the evidence from before the debt crisis is inconsistent with this explanation and that her findings are equally consistent with an explanation based on dynamic inconsistency whereby countries that have less discipline in their borrowing also have less discipline in their monetary policy.

[2] A number of papers have shown how countries with independent central banks have lower inflation (for example, Alesina, 1988, and Cukierman, 1992. Cukierman, Edwards, and Tabellini (1992) show that politically unstable countries have higher inflation and point out that this may reflect the smaller importance of building reputations when time horizons are shorter. Our paper is also related to work on institutional complementarities (for example, Hall & Soskice, 2001).

$\frac{1}{2}\pi^2$. Assume the tax that each individual pays goes to support the benefit system ($T = ub$), and call the replacement rate $r = b/W$. Then the first two terms of V can be expressed in terms of r as $S(r, u) = \log W + (1 - u) \log (1 - ur) + u \log [r(1 - u)]$.

Following Barro and Gordon (1983), the unemployment rate equals the equilibrium (or "natural") rate plus a term that depends negatively on unexpected inflation, $\pi - \pi^e$. The parameter $\beta > 0$ is the slope of the Phillips curve. This assumption is often justified in terms of an imperfect-information story (as in Lucas's model) or in terms of staggered wage or price setting and costly price adjustments. The equilibrium rate, in turn, is made up of a small positive number (which we can call the frictional unemployment), u^f, and a term that depends on the generosity of unemployment benefits (which we can call the structural unemployment), αr. The parameter α denotes the severity of *incentive effects*, whereby higher benefits imply higher unemployment, as in standard search models or efficiency wage models.[3] This is summarized by

$$u = u^f + \alpha r - \beta(\pi - \pi^e). \tag{1}$$

The monetary authority seeks to set inflation in order to maximize aggregate ex post welfare in this economy, taking as given the level of ex ante inflationary expectations of the workers, π^e. The policymaker's inflation-setting problem is

$$\text{Maximize}_\pi \quad V = S(r, u) - \frac{1}{2}\pi^2$$

$$\text{such that} \quad u = u^f + \alpha r - \beta(\pi - \pi^e). \tag{2}$$

In the equilibrium of the one-period game with no uncertainty, expectations must be fulfilled ($\pi = \pi^e$), so that

$$\pi = -\beta \frac{\partial S}{\partial u}, \tag{3}$$

where

$$\frac{\partial S}{\partial u} = -\frac{r(1 - u)}{1 - ur} - \log\left(\frac{1}{r - u}\right) - \frac{u}{1 - u} + \log (1 - u), \tag{4}$$

which is negative for $0 \le r \le 1$.

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[3] The following example illustrates. Assume that a firm pays n workers the wage w to produce output at price p, and that competition ensures zero profits: $\pi = q(n)p - nw = 0$. There are $n = (1 - U)L$ workers, where U is the unemployment rate and L is the labor force. Hence $\pi = \pi(W, U)$, where $W = w/p$. Assume that workers can shirk on their job, in which case effort equals 0, but if caught, they are fired. The expected income from being fired equals the probability of staying unemployed [$= a(U)$, where $\partial a/\partial U > 0$] multiplied by the level of benefits, plus the probability of finding a new job [$= 1 - a(U)$] multiplied by the wage net of taxes and effort costs. The *no-shirking condition* (NSC) equates the value from exerting effort on the job to the value of shirking: $C(W, r, U) = 0$ (where $\partial C/\partial W > 0$, $\partial C/\partial r < 0$, $\partial C/\partial U > 0$). The equilibrium unemployment u is determined by equating the NSC to the zero-profit condition: $u = f(r)$, where $du/dr > 0$. For the dynamic case, see Shapiro and Stiglitz (1984).

Theorem. More generous benefits decrease equilibrium inflation when there are small adverse incentive effects of benefits on unemployment.

Proof. Use equation (3) to calculate

$$\frac{d\pi}{dr} = -\frac{\beta(1 - r)}{(1 - ur)^2 r} + \alpha\beta\left\{\frac{2 - u}{(1 - u)^2} - \frac{r[2 - r(1 + u)]}{(1 - ur)^2}\right\}. \tag{5}$$

When $\alpha \to 0$,

$$\frac{d\pi}{dr} \to -\frac{\beta(1 - r)}{(1 - ur)^2 r} < 0. \tag{6}$$

The welfare state can be expected to affect the equilibrium rate of inflation in two ways. First, there is the direct effect of unemployment benefits on inflation that occurs through the first argument of $S(r, u^f + \alpha r)$. Benefits reduce the individual costs of falling unemployed. In the extreme case, we can imagine the policymaker to become greatly distressed when unemployment occurs at zero benefits. This suggests that $S(0, u) \to -\infty$, and that $\partial S/\partial r > 0$ for low values of r. It is also possible that for very high values of r, higher unemployment benefits may make unemployment more costly to the policymaker. This will occur if society perceives that unemployment benefits replace a large part of the wage and that there is widespread abuse of the system, or in a society that is very much concerned about the tax costs of the welfare state. In such cases, the slope of $S(r, u)$ when r is close to 1 may indeed be negative.

Second, there is an indirect effect of unemployment benefits on inflation that occurs through the natural rate of unemployment. More generous benefits can be expected to increase equilibrium inflation, because they drive up the unemployment rate and introduce a gain from generating unexpected inflation. This, in turn, depends on the severity of incentive effects. For example, in a standard efficiency wage model higher unemployment benefits lead workers to be more prone to shirking at given wages. The quasi labor supply (no-shirking condition) shifts up in the wage-employment plane, and equilibrium unemployment increases (see, for example, Shapiro & Stiglitz, 1984). A similar effect of benefits occurs in a standard search model (see, for example, Lippman & McCall, 1979). The more severe these problems are, the higher the unemployment rate at any given level of unemployment benefits, and the higher the equilibrium rate of inflation.

Summarizing these effects, equation (5) implies that we can expect to observe a negative correlation between the equilibrium rate of inflation and unemployment benefits only when the first term dominates the second, which for the above example is always the case when incentive effects are small. When incentive effects are large (or if different

functional forms are assumed for the utility function or the incentive effects of benefits), this prediction may no longer hold.

B. Empirical Strategy

As a first approach, we estimate a linear reduced form of equation (3) given by

$$INFLATION_{it} = \phi \ BENEFITS_{it} + \delta\Omega_{it} + \eta_i + \lambda_t + \varepsilon_{it},$$
$$(7)$$

where $INFLATION_{it}$ is the rate of inflation in year t in country i, η_i is a country fixed effect, λ_t is a year fixed effect, and ε_{it} is an identically, independently distributed (i.i.d.) error term. Some specifications also add country-specific time trends. Ω_{it} includes explanatory variables that may be expected to affect the equilibrium rate of inflation. One such variable identified in the literature (see Romer, 1993) is OPENNESS, which measures the degree of openness in the economy by dividing total imports by GDP. Another is the extent to which the political preferences in a country lean toward the right. The ideological position of the government may not only affect inflation but may also be correlated with benefit generosity (see Alesina, 1987). For this reason we include the explanatory variable, RIGHT WING, which is similar to those measures used by political scientists to indicate the left versus right position of a government. It is constructed in two steps (see, for example, Hicks & Swank, 1992). In the first step, we collect the number of votes received by each party participating in the cabinet and express it as a percentage of the total votes received by all parties with cabinet representation. In the second step, this percentage of support is multiplied by a left-right political scale (from Castles & Mair, 1984) and summed across all parties to give a continuous variable.

For a measure of the generosity of the welfare state we use the parameters of the unemployment benefit system, BENEFITS. The OECD recently produced this index, partly in response to criticisms of previously available measures of unemployment benefits (see Atkinson & Micklewright, 1990). It is defined as the index of (pretax) unemployment benefit entitlements divided by the corresponding wage (calculated for odd-numbered years by the OECD and linearly interpolated to obtain observations for even-numbered years). This summary measure estimates the situation of a representative individual. Because individuals are entitled to different welfare payments according to their personal situation, the OECD allows for differences in the length of their unemployment spell, in their family situation, and in their previous earnings levels. Specifically, it calculates the unweighted mean of 18 numbers based on all combinations of the following scenarios: (i) three unemployment durations—the first year, the second and third years, and the fourth and fifth years in the unemployment

spell[4]; (ii) three family situations—a single person, a married person with a dependent spouse, and a married person with a spouse in work; (iii) two different levels of previous earnings—average earnings and two-thirds of average earnings. Figure A1 in appendix A shows the cross-sectional relationship between inflation and benefits for the 20 countries in our sample, using decade averages of both variables.

We also report regressions where the benefit data are separated into the two primary dimensions along which unemployment benefit generosity may differ across nations and over time: level and duration. This was done by obtaining from the OECD the original data set used to create the summary measure of benefit entitlements described above.[5] The data were then separated into a measure of the generosity of the replacement rate during the first year (BENEFIT REPLACEMENT) and a measure of the length of time an unemployed worker receives unemployment benefits at an unchanged level (BENEFIT DURATION). Specifically, BENEFIT REPLACEMENT is defined as the OECD index of (pretax) unemployment insurance benefit entitlements (divided by the wage) for an unemployment duration of less than 1 year, calculated as the unweighted mean of six numbers based on all combinations of the following scenarios: (i) three family situations—a single person, a married person with a dependent spouse, and a married person with a spouse in work; (ii) two different levels of previous earnings—average earnings and two-thirds of average earnings. BENEFIT DURATION is an index that captures the length of time the first-year level of benefit generosity is paid and varies on a continuous scale between 1 and 3. It equals 1 if benefits last for just 1 year, 2 if the initial level of benefits is paid out for a period lasting between 1 and 3 years, and 3 if the initial level is paid for a period lasting between 4 and 5 years.[6] If, for example, the initial level of the benefit replacement rate halves after 1 year and then becomes 0 in the fourth and fifth years, BENEFIT DURATION equals 1.5. We obtained data based on consistent definitions for the two dimensions for the 32-year period between 1961 and 1992, except for Portugal (only since 1982) and France (only since 1983), for a total of 595 observations (595 = 32 × 18 + 10 + 9). See the OECD Jobs Study (1994). Appendix B has definitions of all the variables used in our regressions.

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[4] For individuals with a long record of previous employment in the three cases.

[5] We thank David Grubb and Pascal Marianna at the OECD for providing us with the data and for many explanations regarding their construction.

[6] The calculation is as follows: The first-year average benefit replacement rate is subtracted from the second + third-year average replacement rate, and this difference is then divided by the first-year replacement rate to give a number, d_1. Next, the first-year average replacement rate is subtracted from the fourth + fifth-year average replacement rate, and this difference is then divided by the first-year replacement rate to give d_2. The index used is defined as the sum $d_1 + d_2 + 3$. Further details are contained in Appendix B, which also has definitions of all other variables used in the regressions.

TABLE 1.—SUMMARY STATISTICS: 20 OECD COUNTRIES, ANNUAL DATA FOR 1961–1992

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
INFLATION	595	0.065	0.045	−0.007	0.288
BENEFITS	595	0.237	0.136	0	0.632
REPLACEMENT	595	0.393	0.205	0	0.888
DURATION	595	1.805	0.711	1	3.084
UNEMPLOYMENT	595	0.048	0.039	0	0.214
OPENNESS	595	0.309	0.156	0.049	0.929
RIGHT WING	329	5.115	1.594	2.300	7.800

When the benefit data are separated into level and duration measures, the estimated regressions substitute ϕ *BENEFITS$_{it}$* for ϕ^1 *BENEFIT DURATION$_{it}$* + ϕ^2 *BENEFIT REPLACEMENT$_{it}$* in equation (7). Constructing measures of the different dimensions of benefit generosity (*REPLACEMENT* and *DURATION*) is important for two reasons. First, theoretically policymakers may react more to changes in the probability of having unemployed people on zero income (changes in *DURATION*) than to changes in the amount of income they receive (changes in *REPLACEMENT*). Second, the definition of *BENEFIT DURATION* ensures that it is independent of wages; for the level of average earnings (as well as two-thirds of average earnings) for which the various benefit scenarios are calculated appears in both the numerator and the denominator of the two ratios used to calculate this index. The level of earnings therefore cancels out. Consequently any negative effects of inflation on the generosity of benefits relative to wages that may be present (unless the benefit system is fully indexed) should not influence this variable. Table 1 presents summary statistics.

In contrast, the benefit replacement rate may decline in high-inflation periods if benefits are not fully indexed (that is, wages may rise more than benefits). In fact this was not the case in our sample. Not only did the average level of *BENEFIT REPLACEMENT* across all the countries not decline during the period of highest inflation in the sample between 1973 and 1982—on the contrary, this was the period of its biggest increase. However, because it may still be argued that nonindexation of unemployment benefits could be a factor influencing measurements of benefit replacement rates, it is important to use another measure of benefit generosity that does not depend on the nominal levels of benefits and wages (namely, *BENEFIT DURATION*).

We also obtained direct evidence on how benefits are actually set in different countries. Weekly benefits to the unemployed are expressed in most countries as a percentage of wages during a recent period (so that when nominal wages rise due to inflation, benefits are automatically increased to keep the replacement rate constant). However some countries, such as New Zealand, pay a flat level of unemployment benefit with no explicit legislative allowance for adjustments to take account of price changes. *BENEFIT REPLACEMENT* rose from 0.27 to 0.34 in New Zealand over the period 1979 to 1982 when inflation was at its

highest levels due to the oil shock. *BENEFIT DURATION* remained at the constant value of 3 over the whole sample period, including the oil shock years (that is, benefits continued to be paid in all future periods at the first-year level of benefit generosity). In Ireland a flat level of unemployment benefits is paid, as well as graduated supplements equal to a specified percentage of average earnings. Both the *REPLACEMENT* and *DURATION* measures of benefit generosity rose in Ireland over the period 1979 to 1982, the former from 0.52 to 0.53, and the latter from 1.6 to 1.7. The United Kingdom pays flat level unemployment benefits, rather than graduated benefits varying with past wages. In the United Kingdom *BENEFIT REPLACEMENT* decreased from 0.34 to 0.29 between 1979 and 1982, whereas *BENEFIT DURATION* increased, from 2.1 to 2.4.[7] Table 2 shows the average level of inflation and different measures of benefit generosity for each country across our sample.

More generally, there are other institutional variables that have also been used in the prior literature to help explain variations in inflation. For example, Cukierman et al. (1992) use cross-sectional data from 79 countries to find a positive correlation between inflation and political instability, whereas Cukierman (1992) documents a negative relation between inflation and central-bank independence. Cukierman and Lippi (1999) test for the existence of interaction effects between the degree of centralization of wage bargaining and central-bank independence on the level of inflation. They use measures of the degree of the centralization of wage bargaining for the OECD on a discrete 1-to-3 scale measured at three points in time. (Among the 51 country-year observations in their data set, there are a total of eight changes in which countries altered the level of centralization of their wage bargaining systems by one level.) Romer and Romer (1998) emphasize the connection between inflation and the well-being of the poor (see also Albanesi, 2001). Alesina and Roubini (1992) provide evidence that partisan political effects on inflation depend on the form of government (proportional versus majoritarian). Because the available data on most of these variables have low time variability within each of our OECD countries, identification of their effect in the regression specifications

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[7] The principal features of unemployment benefit systems in over 120 nations are provided by *Social Security Programs Around the World*, a U.S. Department of Health and Human Services publication.

TABLE 2.—DESCRIPTION OF DATA: COUNTRY AVERAGES FOR 1961–1992

Country	INFLATION	BENEFITS	REPLACEMENT	DURATION
Australia	0.067	0.208	0.205	3.040
Austria	0.044	0.183	0.262	1.858
Belgium	0.049	0.417	0.464	2.725
Canada	0.055	0.257	0.512	1.523
Denmark	0.069	0.460	0.663	2.064
Finland	0.075	0.228	0.367	1.746
France	0.038	0.366	0.594	1.850
Germany	0.035	0.292	0.400	2.189
Ireland	0.084	0.242	0.366	2.093
Italy	0.090	0.018	0.053	1.000
Japan	0.054	0.109	0.328	1.000
Netherlands	0.046	0.397	0.622	1.851
Norway	0.066	0.189	0.326	1.519
New Zealand	0.086	0.311	0.306	3.053
Spain	0.100	0.226	0.574	1.240
Sweden	0.069	0.218	0.613	1.049
Switzerland	0.040	0.114	0.340	1.004
United Kingdom	0.079	0.231	0.301	2.331
United States	0.050	0.116	0.260	1.334
Portugal	0.132	0.101	0.230	1.959

that we are using (with a full set of fixed effects as controls) is not possible.

There seem to be two views on the appropriate time horizon to analyze problems of dynamic inconsistency. On the one hand, Romer (1993) focuses on the long run, taking a large cross section of countries where inflation and openness are averaged out over approximately 12 years.[8] On the other hand, a number of economists have used Kydland and Prescott (1977) to analyze the short-run profile of inflation. This is the case of Barro and Gordon (1983) and, in particular, Alesina (1988), who exploits the preferences of left- and right-wing governments to predict macro movements following changes in government. In his review of Blinder (1998), Sargent (1999) cites work by Parkin (1993) and Ireland (1999), who both argued that the U.S. time series on inflation can be explained with a version of Kydland and Prescott's time-consistent equilibrium. In the actual models, there is nothing that suggests that these effects take very long. On the contrary, any change in the costs of inflation in the welfare function [equation (2)] is translated instantaneously into changes of the equilibrium inflation rate. Even if one had a view that these effects are only relevant in the long term, the fact that the correlation could be present at business cycle frequency is worth pointing out.

We are agnostic on the appropriate time frame for these forces to affect inflation levels. Thus, we present our analysis both using a yearly panel and collapsing the data into four different time periods (given that our data cover almost four decades). These are 1961–1968, 1969–1976, 1977–1984, and 1985–1992. Each period is 8 years long. Note that this aggregation is more likely to affect the size of the coefficients of those variables that have considerable intra-decade variation. Thus, we expect a larger difference between the annual and decadal estimates of, say, the unem-

[8] The period analyzed starts in 1973, and the inflation data are taken from a 1986 publication.

ployment rate than of benefits. Table 3 shows summary statistics for the period averages across each decade for all our variables. Table 4 shows the correlation coefficients between our variables of interest.

III. Empirical Results

A. Basic Results

Table 5 reports our basic set of results using yearly observations. Regression (1) regresses the rate of inflation on benefits (the OECD summary measure of the parameters of the unemployment benefit system), controlling for country and year fixed effects. The coefficient on BENEFITS is negative and significant at the 1% level. A 1-standard-deviation decrease in BENEFITS (equal to 0.14) is predicted to increase inflation by 0.7 percentage points. This represents 15% of the standard deviation in inflation (which is equal to 0.045).

Regression (2) reports results dividing BENEFITS into separate measures of the two primary dimensions along which unemployment benefit generosity may differ across nations and time, the REPLACEMENT rate and DURATION. Theory predicts that policymakers will be particularly sensitive to duration, because shorter duration makes it more likely that some unemployed individuals have exhausted their unemployment insurance payments. It reports that the effect of the REPLACEMENT rate is insignificant and that DURATION is negative and significant at the 1% level in our base regression that controls for country and year fixed effects. The coefficient on DURATION is −0.02, so a 1-standard-deviation decrease in this variable (equal to 0.71) is predicted to add 1.4 percentage points to inflation. This represents 31% of the standard deviation in inflation. To get another idea of the size of this effect, consider a drop in the

TABLE 3.—SUMMARY STATISTICS: 20 OECD COUNTRIES, FOUR 8-YEAR AVERAGES FOR 1961–1992

Period and Variable	Obs.	Mean	Std. Dev.	Min.	Max.
1961–1968:					
INFLATION	18	0.036	0.012	0.020	0.060
BENEFITS	18	0.165	0.108	0.029	0.358
REPLACEMENT	18	0.278	0.117	0.087	0.484
DURATION	18	1.710	0.793	1.000	3.040
UNEMPLOYMENT	18	0.022	0.015	0.001	0.050
OPENNESS	18	0.230	0.112	0.054	0.499
RIGHT WING					
1969–1976:					
INFLATION	18	0.085	0.021	0.051	0.126
BENEFITS	18	0.209	0.129	0.018	0.496
REPLACEMENT	18	0.353	0.182	0.053	0.737
DURATION	18	1.749	0.698	1.000	3.040
UNEMPLOYMENT	18	0.028	0.018	0.003	0.062
OPENNESS	18	0.296	0.147	0.073	0.668
RIGHT WING	16	5.027	1.170	2.900	7.100
1977–1984:					
INFLATION	20	0.098	0.048	0.035	0.234
BENEFITS	20	0.264	0.135	0.008	0.564
REPLACEMENT	20	0.440	0.208	0.024	0.831
DURATION	20	1.825	0.685	1.000	3.052
UNEMPLOYMENT	20	0.063	0.032	0.014	0.124
OPENNESS	20	0.322	0.156	0.086	0.747
RIGHT WING	17	5.272	0.921	3.090	7.003
1985–1992:					
INFLATION	20	0.047	0.025	0.017	0.121
BENEFITS	20	0.299	0.132	0.024	0.589
REPLACEMENT	20	0.487	0.211	0.071	0.867
DURATION	20	1.905	0.649	1.000	3.082
UNEMPLOYMENT	20	0.075	0.041	0.023	0.186
OPENNESS	20	0.378	0.178	0.116	0.863
RIGHT WING	17	5.289	1.473	2.900	7.800

TABLE 4.—CORRELATION COEFFICIENTS: 20 OECD COUNTRIES, FOUR 8-YEAR AVERAGES FOR 1961–1992

Variable	INFLATION	BENEFITS	REPLACEMENT	DURATION	UNEMPLOYMENT	OPENNESS
BENEFITS	−0.068	1				
REPLACEMENT	0.027	0.787	1			
DURATION	−0.086	0.572	0.029	1		
UNEMPLOYMENT	0.190	0.358	0.325	0.124	1	
OPENNESS	0.047	0.550	0.368	0.315	0.290	1
RIGHT WING	0.008	0.036	−0.031	0.053	0.307	0.031

DURATION index variable of 2 units (that is, a shift from benefits lasting 3–4 years to less than 1 year), equivalent to a move from Australia at the top of the sample to Italy (which has the least generous unemployment insurance system in our sample). This is predicted to increase inflation by 4.1 percentage points. Mean inflation in Australia over the sample period equaled 6.7%, while in Italy it equaled 9%. Hence the effect of DURATION can account for the different inflation experiences in these two countries.

The previously reported set of results can be thought of as producing reduced-form estimates of the effect of benefits on inflation, based on the first-order condition in equation (3). The coefficient on our benefit explanatory

variable captures several effects.[9] These can be seen by examining equation (3) above, which identifies a direct effect of benefits on equilibrium inflation that occurs through the cost of unemployment (via the cost of risk and taxes) as well as an indirect effect that occurs through the natural rate of unemployment. Higher benefits may increase the natural rate due to adverse incentive problems (where $u^n = u^f + \alpha r$). In order to obtain an estimate of the direct effect, regression (3) reports results controlling for the unemployment rate, UNEMPLOY-

[9] Just as in Romer (1993), where the coefficient on openness captures a number of channels through which openness affects inflation.

TABLE 5.—INFLATION REGRESSIONS USING THE OECD MEASURE OF BENEFIT ENTITLEMENTS: 20 OECD COUNTRIES, ANNUAL PANEL FOR 1961–1992

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
BENEFITS	−0.051 (0.017)							
REPLACEMENT		0.006 (0.010)	0.006 (0.010)	0.009 (0.010)	−0.035 (0.014)	0.013 (0.015)	0.006 (0.027)	0.006 (0.018)
DURATION		−0.020 (0.003)	−0.022 (0.003)	−0.023 (0.003)	−0.025 (0.005)	−0.021 (0.006)	−0.022 (0.007)	−0.022 (0.006)
UNEMPLOYMENT			−0.226 (0.057)	−0.189 (0.057)	−0.624 (0.076)	−0.536 (0.100)	−0.226 (0.151)	−0.226 (0.097)
OPENNESS				−0.105 (0.026)	−0.191 (0.039)			
RIGHT WING					−0.001 (0.001)			
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country-specific year trends						Yes		
No. of observations	595	595	595	595	329	595	595	595
Adj. R^2	0.69	0.70	0.71	0.72	0.82	0.77	0.71	0.71

NOTE. Dependent variable in all regressions is the inflation rate. Standard errors in parentheses. Standard errors in regression (7) are clustered by country. Standard errors in regression (8) are clustered by country-decade combination. All regressions include country and year fixed effects. Regression (6) also includes country-specific time (year) trends.

MENT.[10] Note that this proxy is suitable to the extent that inflationary expectations are being equated to actual inflation outcomes from year to year [see equation (1) with $\pi = \pi^e$]. To the extent that expectations are not being fulfilled yearly, we may not identify the long-run equilibrium relationship between inflation and the natural rate of unemployment (which our theory from section II predicts is positive), but instead identify a short-run Phillips curve, when estimating regressions based on annual data. Because the coefficient on UNEMPLOY-MENT in regression (3) is negative and significant, this indicates that inflationary expectations may be taking longer than 1 year to adjust. Controlling for the unemployment rate does not change the size and significance of the coefficient on DURATION. (In the next section, where we estimate regressions using a decadal panel, the correlation between inflation and unemployment becomes insignificant across all specifications.)

Regression (4) adds OPENNESS as a control, as suggested by Romer (1993). The coefficient on DURATION retains its size and significance. OPENNESS has a negative and well-defined effect on inflation. This contrasts with Romer's insignificant cross-sectional estimates for the OECD. It is interesting to note that a 1-standard-deviation increase in openness (equal to 0.16) brings about the same reduction in inflation rates as a 1-standard-deviation increase in our measure of the generosity of the welfare state (in both cases just under 1.7 percentage points).

[10] The previous version of this paper reports similar results using a more elaborate measure of the natural rate of unemployment obtained by a cubic interpolation of the unemployment series for each country [as suggested by Staiger, Stock, and Watson (1997)]. Another way of obtaining a time-varying natural rate of unemployment for each country is by calculating a moving average of the time series of unemployment rates (see Mankiw, 1997). We also tried this method, and our main results were unaffected.

Regression (5) reports results that control for RIGHT WING, which is an index of left versus right political party strength. This explanatory variable was included to control for the possibility of omitted variable bias in that more right-wing political parties may have policies that affect both the level of inflation and the unemployment benefits. Presumably, right-wing parties tend to prefer lower inflation rates and lower benefits than do left-wing parties. See, for example, Alesina and Roubini (1992). Then the omission of a variable measuring political ideology should have biased our results away from finding the negative relationship that our previous regressions have identified. Hence the true effect may have been underestimated. We constructed the variable on right-wing politics using data on the electoral performance of the various parties with cabinet representation (see the data definitions and sources in appendix B). Because the data were not available for a number of countries, our sample is smaller in this regression. Still, the results are informative. The coefficient on DURA-TION is negative and significant at the 1% level. Its size is again very similar to that of previous regressions. The coefficient on UNEMPLOYMENT is also negative and significant, although its absolute size is (more than 3 times) larger than that of the previous regression. A similar result occurs for OPENNESS, whose coefficient is 82% larger than that in regression (4). Finally, the effect of RIGHT WING is negative but only significant at the 16% level. In this sample, and to the extent that the estimated effects are taken as causal, a more generous welfare state is a more reliable way to get low inflation than having a right-wing government.

We have also experimented with breaking the sample into different time periods. If, for example, we divide our time period into halves and estimate regression (3) from table 5 over a pre- and a post-1975 period, the coefficients on

TABLE 6.—INFLATION REGRESSIONS USING THE OECD MEASURE OF BENEFIT ENTITLEMENTS BY DECADES:
20 OECD COUNTRIES FOR FOUR 8-YEAR PERIODS BETWEEN 1961–1992

Variable:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
BENEFITS	−0.057							
	(0.038)							
REPLACEMENT		0.009	0.008	0.014	0.014	0.039	0.008	0.008
		(0.024)	(0.025)	(0.024)	(0.031)	(0.059)	(0.037)	(0.029)
DURATION		−0.023	−0.023	−0.024	−0.020	−0.030	−0.023	−0.023
		(0.007)	(0.007)	(0.008)	(0.012)	(0.017)	(0.008)	(0.006)
UNEMPLOYMENT			−0.040	0.031	−0.328	−0.225	−0.040	−0.040
			(0.142)	(0.145)	(0.217)	(0.342)	(0.203)	(0.161)
OPENNESS				−0.121	−0.066			
				(0.055)	(0.069)			
RIGHT WING					−0.003			
					(0.002)			
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Decade dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country-specific decade trends						Yes		
No. of observations	76	76	76	76	50	76	76	76
Adj. R^2	0.84	0.85	0.85	0.86	0.90	0.91	0.85	0.85

NOTE. Data averaged at the decade level. Dependent variable in all regressions is the inflation rate. Standard errors in parentheses. Standard errors in regression (7) are clustered by country. Standard errors in regression (8) are clustered by country–two-decade combinations. (The two-decade periods are 1961–1976 and 1977–1992.) All regressions include country and year fixed effects. Regression (6) also includes country-specific time (decade) trends.

DURATION now equal −0.017 (s.e. = 0.007) and −0.083 (s.e. = 0.026), respectively. That is, both remain negative and significant at the 2% level, and the coefficient in the second period is more negative. The results are similar even if we divide the sample into, for example, a pre- and a post-1980 period.

B.   Serial Correlation and Analysis with Decadal Panel

A potential problem in our analysis is that the error term is serially correlated. The inclusion of country fixed effects controls for a particular kind of serial correlation, namely the kind that shifts the mean of the error term. There are, of course, a number of economic forces that would induce a different structure of serial correlation. A simple possibility is that there are trended variables in each country that drive both movements in the welfare state and in the inflation rate. In regression (6) we include country-specific time trends, as well as country and year fixed effects. There are consequently 69 control variables in this specification of the error term (19 country dummies + 19 country-specific year trends + 31 year dummies). The coefficient of DURATION is again negative and significant at the 1% level and of a similar size to those in previous regressions.

Another possible solution in the presence of positive serial correlation is to allow for an arbitrary covariance structure within countries over time. A common approach is to introduce a cluster adjustment at the group level that, essentially, multiplies the standard errors by a factor that depends positively on the product of the size of the cluster and the intracluster residual correlation.[11] Regression (7) in

table 5 presents our results when clustering at the country level is allowed. Comparing with the unclustered results [see column (3)], the standard errors on DURATION more than double, whereas those on UNEMPLOYMENT are multiplied by almost three. The coefficient on DURATION is still significant at the 1% level. As pointed out in Angrist and Lavy (2002), inference in this case turns on an asymptotic argument based on the number of clusters. Given that regression (7) only uses 20 clusters, it can be argued that these are too few to provide an accurate approximation to the sampling distribution [those authors cite the work of Thornquist and Anderson (1992)]. Regression (8) uses the 76 decade-country combinations as the cluster unit. The results are largely unchanged.

Finally, we also experimented with including a lagged dependent variable (while clustering using the decade-country combinations). These results are available upon request. To see why this may help, consider the extreme case of an A(1) process for inflation. In that case, including a lagged dependent variable should leave no role for the welfare state in the determination of inflation. The coefficient on UNEMPLOYMENT equals −0.194 (s.e. 0.052), whereas that on DURATION equals −0.008 (s.e. 0.003). Given that the coefficient on lagged inflation is 0.628 (s.e. 0.042), the long-run effect of DURATION is equal to −0.022 [= −0.008/(1 − 0.628)], which is similar to previous estimates.

Table 6 presents the empirical results using a different strategy that involves moving to group-level data. Specifically, the data are collapsed into four time periods of equal

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[11] The variance formula for the calculation of the clustered standard errors is given by: $\hat{\Sigma} = q_c(X'X)^{-1}(\sum_{h=1}^{M} \mathbf{u}_h' \mathbf{u}_h)(X'X)^{-1}$, where $\mathbf{u}_h = \sum_{j \in G_h} \mathbf{u}_j$; $G_1, G_2, \ldots, G_M$ are the clusters; M is the number of clusters; $\mathbf{u}_j = (y_j - \mathbf{x}_j \mathbf{b})\mathbf{x}_j$; and $q_c = [(N-1)/(N-k)]M/(M-1)$; and y_j, $\mathbf{x}_j$, $\mathbf{b}$, X, N, and k follow standard econometric notation (for further details, see STATA (2001), p. 87). This variance estimator coincides with the Huber-White variance estimator when each cluster contains one observation.

length. These are from 1961 to 1968, from 1969 to 1976, from 1977 to 1984, and from 1985 to 1992. Each is 8 years in length. They correspond roughly to the four decades included in our study, leaving 76 observations, as we lose the first two decades for Portugal and France. A further advantage in this case is that one can investigate whether the Barro-Gordon positive theory of inflation based on time inconsistency of policymaking has predictions for high-frequency data (see section II on the empirical strategy). The results of the effects of benefit generosity on inflation, where both variables are now measured at the decade level, are very similar to those presented using the annual panel.

The size of the coefficient of interest (*DURATION*) ranges between -0.02 and -0.03 (taken across all the table 6 specifications). For example, in regression (2) it equals -0.023 (compared to -0.020 in the comparable column in table 5), which is significant at the 1% level. A 1-standard-deviation decrease in this variable (equal to 0.70 measured at the decade level) is predicted to add 1.6 percentage points to inflation. This represents 41% of the standard deviation of the inflation rate (measured across decades). The coefficient on *DURATION* remains significant at the 1% level, even after controlling for *UNEMPLOYMENT* in regression (3) (which is insignificant) and *OPENNESS* in regression (4). If higher benefit durations cause higher unemployment, our theory suggested that the coefficient on benefit duration would have been more negative once the control for unemployment was added. The degree to which this effect occurs depends of the size of the adverse incentive effects of benefits and also the existence of positive effects of the natural rate of unemployment on equilibrium inflation, which have not been detected.

The significance of *DURATION* drops in some cases, notably in regressions (5) and (6), where the coefficient is significant only at the 10% level. Note that in some cases the number of degrees of freedom falls considerably. For example, in regression (5), where *RIGHT WING* data are employed, there are only 50 observations from which to estimate the coefficients of the explanatory variables (which include the decade and country fixed effects). The single most notable effect that moving to decade-long averages has on the results is on the coefficient of *UNEMPLOYMENT*. Because we are now no longer measuring short-run effects, we may expect to lose the negative (short-run Phillips curve) relationship between inflation and unemployment that appeared in the previous table (which used yearly data). Instead we may expect to now see a vertical (long-run) Phillips curve. The results are consistent with this view. Table 6 shows no evidence of a significant negative trade-off between inflation and unemployment using decade-long averages, as the coefficient on unemployment is insignificant across all specifications.

Regressions (7) and (8) control for any remaining serial correlation (after already taking decade averages) by allowing for clustering at the country level and the country–two-

decade level, respectively. [The two-decade periods used for regression (8) are 1961–1976 and 1977–1992.]

C. Simultaneity and the Endogenous Welfare State

Traditionally, economists have taken labor market institutions as exogenous. This is true both in the macroeconomics literature we referred to above, and the literature in labor economics that studies the effect of unemployment benefits on unemployment rates [for example, Layard, Nickell, and Jackman (1990) and Blanchard and Wolfers (1999)]. It can be argued, however, that basic economic forces affect benefits. These include a desire to balance insurance and tax considerations, or to reduce the adverse incentive effects of benefits [see the models of Wright (1986), Atkinson (1990), Saint Paul (1996), Hassler et al. (1999), Rodrik (1998), and Di Tella and MacCulloch (1995, 2002), as well as the empirical evidence in the last three of these]. For example, Di Tella and MacCulloch (1995) show that employed individuals in high-unemployment regions in the United Kingdom are more likely to think that unemployment benefits are too low than are those living in low-unemployment regions. Furthermore, unemployment benefits exhibit considerable variation across countries and over time, particularly at times of economic turbulence (such as the oil shocks in the 1970s).[12] Unfortunately, a convincing instrument in this setting is not available, so a proper investigation into these issues has to be left for future research.

In terms of our model, it is natural to think of benefit institutions as being fixed prior to the setting of inflation. This will be the case when benefits are costly to change, due, for example, to a protracted legislative process. Two simple points are worth mentioning. First, consider a two-stage game where benefits are chosen in the first stage to maximize V, and in the second stage the monetary authority sets inflation [according to the problem (2)]. Solving backward, we obtain $\pi = \pi(r, \alpha, u^f)$ [from equation (3)]. In the absence of concerns about the effect of benefits on inflation,

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[12] The two periods that include the oil price shocks in 1973–1975 and 1979–1981 precede and include the two periods that include the largest increases in U.S. benefit generosity. Di Tella and MacCulloch (2002) provide direct evidence on how negative shocks result in longer benefit durations by examining the legal environment defining benefit provision. Countries as diverse as Canada, Japan, and the United States have laws actually *stating* that benefits depend on aggregate unemployment conditions. In the United States, the Federal/State Extended Compensation Act of 1970 established a second layer of benefits for claimants who exhaust their regular entitlement during periods of relatively high unemployment in a state. This program provides for up to 13 extra weeks of benefits at the claimant's usual weekly benefit amount. The benefits are triggered if the state's insured unemployment rate for the past 13-week period is 20% higher than the rate for the corresponding period in the past two years and the rate is at least 5%. Extended benefits cease to become available when the insured unemployment rate does not meet either the 20% requirement or the 5% requirement. In 1973 the 13-week rule applied, but in 1975, as labor market conditions worsened in the face of the first oil shock, federal law made unemployment insurance payable "for additional 26 weeks in cases of high unemployment." This applied until 1983, when federal law reduced the extension back to 13 weeks. Hence, in the United States, adverse shocks that change the unemployment rate also change benefit durations.

let the optimal replacement rate equal $r^* = r(\alpha, u^f)$. A standard result from this type of problem is that higher disincentive effects (that is, large α) should optimally imply lower benefits. But once the effect of benefits on equilibrium inflation is taken into account, $dV/dr|_{r=r^*} = -\pi(d\pi/dr)$. Thus, $dV/dr|_{r=r^*} > 0$ when higher benefits decrease the level of equilibrium inflation. In other words, benefits should be set at a higher level than they would have been in the absence of their impact on monetary policy.

Second, the structure of the problem suggests there may be a role for inflows as an instrument. To see this, note that, keeping the level of unemployment constant, the level of inflows into the unemployment pool drives changes in the optimal level of unemployment benefits. As long as the lag in determining benefits is longer than the lag in setting inflation, the level of inflows can be thought of as affecting the level of inflation only through its effect on the desired level of benefits, particularly if the level of unemployment is kept constant. For example, if inflation is set each period but benefits are set to maximize a T-period discounted sum of expected future values of V, then higher inflows at $T = 0$ affect optimal benefit setting but only affect inflation to the extent they translate into higher future rates of unemployment. Thus a priority for future research is to obtain good data on inflows.

We experimented by instrumenting benefit duration with its lagged value as well as the change in the unemployment rate (as a rough proxy for the inflow rate). The coefficient on duration remains negative and significant (equal to -0.019; s.e. $= 0.004$). In the first-stage regression, positive changes in unemployment increase the level of benefit duration.

A point in connection with the previous literature is worth making. Romer (1993) identifies openness as an independent determinant of the inflation rate. He argues that there are fewer output gains due to surprise inflation in more open economies, particularly when there is a floating exchange rate regime. But, as emphasized above, previous work has found evidence that unemployment risk affects the level of benefit generosity in nations [see Di Tella and MacCulloch (1995, 2002) and Rodrik (1998)]. Inasmuch as it can be argued that a basic proxy for unemployment risk is the degree of openness in the economy, our paper suggests that it is possible that the hypothesized effect of openness on inflation identified by Romer (1993) is operating through unemployment benefits. We conducted some tests and found that the part of OPENNESS that covaries with DURATION is quite large (correlation coefficient $= 0.31$), suggesting that the precise channel through which OPENNESS affects the inflation rate is an open question.

A potential solution for identifying the relative importance of the two channels would be to run a 2SLS regression where benefits were instrumented with the inflow rate into unemployment (if these data were available), also controlling for openness. If variations in risk are fully captured by changes in the inflow rate, then the coefficient on openness

should only be significant to the extent that there exists another channel through which it affects equilibrium inflation.

VI. Conclusions

An influential approach in macroeconomics views the rate of inflation as the outcome of an unresolved dynamic inconsistency problem. Since the first paper by Kydland and Prescott in 1977, a number of papers have developed the approach further to allow for such factors as asymmetric information, stochastic environments, and many countries. The approach has become standard in macroeconomics and is described in most textbooks in the subject. There is, however, remarkably little empirical evidence that can be used to shed light on the relevance of this approach. The paper by Romer (1993) is an exception, as it shows that there is a negative relationship between inflation and openness for a sample of low- and middle-income economies. He finds no evidence of such a relationship holding in the sample of high-income economies: "The results are thus consistent with the view that these countries have largely overcome the dynamic inconsistency of optimal monetary policy" (Romer, 1993, p. 871). In other words, there is no evidence that these types of problems are present precisely in the economies for which the theory was originally developed.

We document a negative relationship between inflation and a measure of the welfare state (unemployment benefits) that should be expected in such a model. Because the costs of unemployment depend on benefit generosity, the welfare gains from the temporary reduction of unemployment due to surprise inflation are also affected by benefits. In other words, building a welfare state has a similar effect to appointing a conservative central banker. We use a newly available OECD data set that measures an unemployed worker's benefit allowance across 18 different types of states, such as marital status and duration of unemployment spell. The data do not depend on the actual number of workers in each state or consequently on the unemployment rate. Using panel data from 1961 to 1992, and controlling for country and time fixed effects, country-specific time trends, and other covariates, inflation is found to depend negatively on the duration of unemployment benefits. The effects are economically significant: a 1-standard deviation decrease in benefit duration is predicted to add 1.4 percentage points onto inflation, or 31% of the standard deviation in inflation. The relationship holds in a panel of OECD countries, a region where Romer (1993) finds no commitment problems.

We point out that unemployment benefits may be endogenously determined. Because a plausible determinant of benefits is the level of openness in the economy, it is an open question through which precise channel openness is affecting the inflation rate.

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APPENDIX A

FIGURE A1.—INFLATION RATE VERSUS BENEFITS (FOUR 8-YEAR AVERAGES): 1961–1992, 20 OECD NATIONS
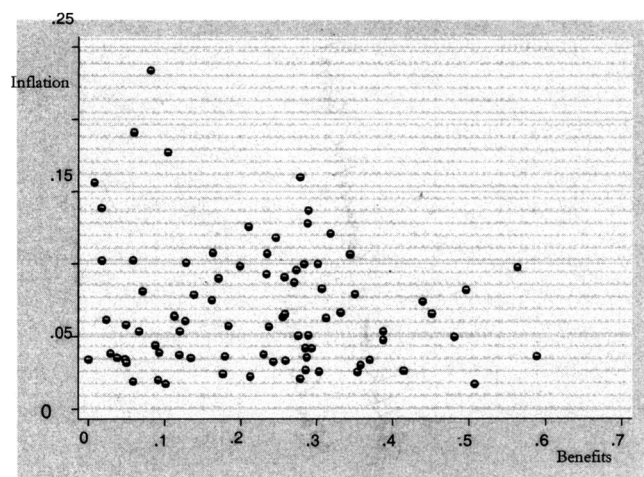
APPENDIX B

Description of Data Set

1. Sample of 20 Countries

Australia, Austria, Belgium, Canada, Denmark, Finland, France (from 1983), Germany, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Portugal (from 1982), Spain, Sweden, Switzerland, United Kingdom, United States.

2. Variable Definitions

INFLATION RATE: Rate of change of GDP deflator from the OECD-CEP data set (1950–1992).

BENEFITS: The OECD summary measure index of (pretax) unemployment insurance benefit entitlements divided by the corresponding wage (calculated for odd-numbered years and linearly interpolated to obtain observations for even-numbered ones). The index estimates the situation of a representative individual. It calculates the unweighted mean of 18 numbers based on all combinations of the following scenarios: (i) Three unemployment durations (for persons with a long record of previous employment): the first year, the second and third years, and the fourth and fifth years of unemployment. (ii) Three family and income situations: a single person, a married person with a dependent spouse, and a married person with a spouse in work. (iii) Two different levels of previous earnings: average earnings and two-thirds of average earnings. See the OECD Jobs Study (1994).

BENEFIT REPLACEMENT: The OECD index of (pretax) unemployment insurance benefit entitlements divided by the wage calculated as the unweighted mean of six numbers based on all combinations of the following scenarios: (i) unemployment duration of less than 1 year. (ii) Three family and income situations: a single person, a married person with a dependent spouse, and a married person with a spouse in work. (iii) Two different levels of previous earnings: average earnings and two-thirds of average earnings. See the OECD Jobs Study (1994).

BENEFIT DURATION: An index capturing the duration of unemployment benefits, calculated as follows. Let R^1 equal the unweighted mean of six measurements that summarize benefit generosity in the first year of unemployment ($= b^{m1}/w$). Let R^2 equal the unweighted mean of 6 measurements that summarize benefit generosity in the second and third years of unemployment ($= b^{m2}/w$). Let R^3 equal the unweighted mean of six measurements that summarize benefit generosity in the fourth and fifth years of unemployment ($= b^{m3}/w$). The six measurements are based on all combinations of the following scenarios: (i) Three family and income situations: single, married with dependent spouse, and married with a spouse in work. (ii) Two different levels of previous earnings: average earnings and two-thirds of average earnings [calculated in the OECD Jobs Study (1994)]. The index, BENEFIT DURATION, is then obtained from the sum $(R^2 - R^1)/R^1 + (R^3 - R^1)/R^1 + 3$. This expression can be simplified to give $(b^{m2} + b^{m3})/b^{m1} + 1$.

UNEMPLOYMENT: The standardized unemployment rate from the OECD-CEP data set.

OPENNESS: Imports divided by GDP from the OECD-CEP data set.

RIGHT WING: Index of left versus right political party strength, defined as the sum of the number of votes received by each party participating in cabinet expressed as a percentage of total votes received by all parties with cabinet representation, multiplied by a left-right political scale constructed by political scientists. Votes are from Mackie and Rose's (1982), International Almanac of Electoral History, cabinet composition is from The Europa Yearbook (1969–1989 editions), and the left-right scale is from Castles and Mair (1984).