A Reduced-form Equation for the Unemployment Rate Estimated from a Panel of Nineteen OECD Countries

DIMITRIS HATZINIKOLAOU*
University of Ioannina, Greece

KONSTANTINA GIRTZIMANI
University of Ioannina, Greece

CHRISTOS MOUSAFIRIS
University of Ioannina, Greece

Abstract: Using a dynamic fixed-effects model for a panel of nineteen OECD countries and annual aggregate data, 1970-2008, we estimate a reduced-form equation for the observed unemployment rate. We find that wage-push factors, such as employment protection and union power, increase the unemployment rate, whereas net immigration reduces it, although this last effect is practically negligible. We also find that the labour market is characterised by strong "lagged adjustment processes," so the governments of the countries in the sample have an additional incentive to fight current unemployment, to prevent it from becoming prolonged.

* Corresponding author: Dimitris Hatzinikolaou, University of Ioannina, Department of Economics, 451 10 Ioannina, Greece.

E-mail: dhatzini@uoi.gr

I INTRODUCTION

In the Organisation for Economic Cooperation and Development (OECD) countries, the unemployment rate was (on average) around 2 per cent in the 1960s, but rose to about 4 per cent in the 1970s, to about 8 per cent in the 1980s, and fluctuated around that level since then (see Blanchard, 2006; Nickell, Nunziata and Ochel, 2005, Table 1). This behaviour of the unemployment rate may reflect the existence of persistence mechanisms through which initial adverse shocks (e.g., the two oil shocks of the 1970s, the productivity slowdown in the 1970s and 1980s, and various policy changes) have lasting effects on unemployment, the hysteresis effect (Blanchard, 2006; Karanassou, Sala and Snower, 2007).

It may also reflect, however, an increase in the "non-accelerating inflation rate of unemployment" (NAIRU) caused by hysteresis (Blanchard and Summers, 1986, p. 15; Tobin, 1980, p. 61) or by changes in labour-market institutions. Examples include union power, employment protection laws (which probably discourage hiring), and social-security benefits, which discourage a worker to search for a job or take a marginal job; see Blanchard (2006); Hatzinikolaou and Kammas (henceforth HK, 2010) and Layard, Nickell and Jackman (1991, pp. 74, 211, 254-258, 509). We believe that these factors were at least partly responsible for the above behaviour of the unemployment rate, since in many OECD countries union power and employment protection were sharply increased in the 1970s and the rates of growth of social-security benefits were high in the 1970s, in the early 1980s and in the early 1990s.

In this paper, we estimate a reduced-form equation for the observed unemployment rate. This equation is derived from a dynamic structural model for the labour market that incorporates the hysteresis effect and the effects of the above institutional variables. Our model differs from that of Karanassou, Sala and Salvador (2008) in that it incorporates some innovative features; e.g., the wage-setting equation includes the growth rate of net immigration and several new wage-push factors, namely, indices of union power and of employ-ment protection and the growth rate of social-security benefits in percent of GDP.[1]

After presenting our estimating equation (Section II), we describe the data and their sources, conduct unit-root tests (Section III), and estimate the equation using a dynamic fixed-effects model for a panel of nineteen OECD countries for the time period 1970-2008 (Section IV). We find that wage-push

[1] The structural model is available from the first author upon request. Thanks to the suggestion of an anonymous referee of this journal, for space considerations, we do not present it here, since its differences from standard models (e.g., Layard, Nickell and Jackman, 1991) are minor and since we do not estimate it.

factors, such as employment protection and union power increase the unemployment rate, whereas net immigration reduces it, although the latter effect is practically negligible. We also find that the long-run effect of the capital stock is zero. Finally, we find that strong "lagged adjustment processes" are at work in the labour market. In Section V, we offer some policy conclusions.

II THE ESTIMATING EQUATION

Our estimating equation for the observed unemployment rate is:

$$u_{it} = \delta_0 + \delta_1 k_{it} + \delta_2 k_{i,t-1} + \delta_3 gun_{it} + \delta_4 f_{it} + \delta_5 gsb_{it}$$
$$+ \delta_6 gnm_{it} + \delta_7 t + \delta_8 u_{i,t-1} + \delta_9 u_{i,t-2} + \varepsilon_{it}, \tag{1}$$

where u_{it} = observed unemployment rate of country i in year t; k_{it} = logarithm of firms' observed capital stock; gun_{it} = growth rate of tun_{it}, a proxy for union power; f_{it} = an index of the intensity of firing restrictions; gsb_{it} = growth rate of social-security benefits in percent of GDP; gnm_t = growth rate of net immigration; t = time trend, which serves to capture different deterministic trends in the variables; and ε_{it} = error term. The expected signs (in accordance with our structural model) of the coefficients δ_2, δ_3, δ_4 and δ_5 are all positive, whereas those of δ_0, δ_1, δ_6, δ_7, δ_8 and δ_9 cannot be determined. We expect $\delta_2 > 0$, because we assume a partial-adjustment model for the capital stock. The higher is last year's actual capital stock the lower is this year's *desired* capital stock, which may reflect the current level of productivity, hence the lower is this year's demand for labour and the higher the current unemployment rate. We expect $\delta_3 > 0$, because, as union power grows faster, wages rise, quantity of labour supplied increases and quantity of labour demanded falls, thus causing the unemployment rate to increase. Next, we expect $\delta_4 > 0$, because stricter employment protection measures discourage hiring new workers, encourage labour supply (as jobs become more secured, hence more attractive) and strengthen the already employed workers' bargaining power, thus causing wages to rise. As a result, the unemployment rate rises. Finally, we expect $\delta_5 > 0$, because, as social-security benefits grow faster, the reservation wages of the unemployed relative to their marginal productivity in employment rise (Tobin, 1980, p. 59), which causes the observed wages to rise, thus causing the unemployment rate to rise. The long-run effect of the capital stock on the observed unemployment rate implied by Equation (1) is $\eta = (\delta_1 + \delta_2)/(1 - \delta_8 - \delta_9)$, so the hypothesis that this effect is zero can be stated as H$_0$: $\delta_1 + \delta_2 = 0$.

III THE DATA: DEFINITIONS, SOURCES AND UNIT-ROOT TESTS

The panel consists of the following 19 OECD countries: Australia, Austria, Belgium, Denmark, Finland, France, Ireland, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, the UK, Canada, Switzerland, Germany, Greece and the US. The first 14 of these countries are those in HK (2010), who considered only countries where the index for firing restrictions exhibits substantial variability. We have added the last five countries for two reasons: (1) to avoid possible criticisms of sample-selection bias, by including three countries (Canada, Switzerland and the US) where the index for firing restrictions exhibits no variability at all during the sample period and two more countries (Germany and Greece) where that index exhibits little variability; and (2) to check the robustness of our empirical estimates to substantial changes in the sample by estimating Equation (1) using both the 19- and the 14-country panel, 1970-2008 in both cases.

The sources of the data are as follows: (1) the *OECD Economic Outlook* for the variables u_{it} and gsb_{it}, which are measured as percentages, as in the case of the growth rates defined below (e.g., a 5 per cent unemployment rate is written as 5); (2) the *European Commission*, *Economic and Financial Affairs* (*AMECO*) for k_{it} = net capital stock in constant prices of 2005 in common currency (billions of euros); (3) the *World Bank* for gnm_{it} = growth rate of net legal migration (number of legal immigrants minus number of emigrants);[2] (4) the *OECD Labour Force Statistics* for tun_{it} = the ratio of wage and salary earners that are trade-union members divided by the total number of wage and salary earners; and (5) the *OECD Employment Protection Index* (ranging from 0 to 6, available from 1985 onward) and the Nickell *et al.* (2001) index (ranging from 0 to 2, available from 1970 to 1995) combined to construct the "employment protection" index f_{it}, which ranges from 0 (least restrictions) to 6 (most restrictions). Note that the series u_{it}, gsb_{it}, f_{it} and gnm_{it} are the same as those used by HK (2010). Note also that HK (2010) estimate the effects of gsb_{it}, f_{it} and gnm_{it} on the NAIRU and not on the observed unemployment rate, u_{it}, so our estimates of these effects are expected to be smaller.

Before estimation, we must test for unit roots. We use the following four tests: IPS_μ and IPS_τ of Im, Pesaran and Shin (2003) and *ADF-Fisher*$_\mu$ and *ADF-Fisher*$_\tau$ (see Maddala and Kim, 1998, p. 137), where the subscripts μ and τ denote level and trend stationarity, respectively. Table 1 reports the results for the 19-country panel, which suggest that all the variables involved in Equation (1) are I(0). The results for the 14-country panel are similar, so, to save space, we do not report them.

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[2] When calculating the growth rate gnm_{it} from the level of net immigration one must ensure that the signs of the observations of gnm_{it} are correct, because net immigration takes both positive and negative values.

Table 1: *Unit-root Tests for the 19-country Panel*

Test Variable	IPS_μ	IPS_τ	Fisher ADF_μ	Fisher ADF_τ	*Decision*
u_t	−4.39***	−1.47*	81.15***	53.15*	I(0)
f_t	−0.95	−2.54***	61.33***	60.14***	I(0)
tun_t	2.01	1.83	43.12	32.27	I(1)
gun_t	−7.45***	−7.70***	131.32***	134.45***	I(0)
k_t	4.93	−3.86***	15.22	88.22***	I(0)
gsb_t	−9.50***	−9.08***	168.78***	155.80***	I(0)
gnm_t	−13.51***	−11.52***	244.39***	190.52***	I(0)

Notes: (1) ***, **, * denote statistical significance at the 1 per cent, 5 per cent and 10 per cent level respectively; (2) the null hypothesis is that the series has a unit root; (3) *IPS* is the Im, Pesaran and Shin (2003) test; (4) the subscript μ indicates that only an intercept is included in the testing regression, whereas τ indicates that both a constant and a trend are included; (5) we do not report the test values for the first differences of the variables, because they are all clearly I(0); (6) the results of this table have been produced by the econometric program *Eviews*; (7) the unit-root test results for the 14-country panel are similar.

IV ECONOMETRIC ANALYSIS AND RESULTS

Using the 19-country panel, we estimate Equation (1) by a fixed-effects method that is robust to autocorrelation and heteroscedasticity. Because lagged values of the dependent variable are used as explanatory variables, our estimates will be consistent only to the extent that the time dimension of our panel (T = 39) is large enough (see Baltagi, 2001, p. 130). The hypothesis $\delta_7 = 0$ could not be rejected (*p*-value = 0.32), so Equation (1) was re-estimated without a time trend. The results are reported in Part A of Table 2.

First, note that the signs of the estimates of $\delta_2, \delta_3, \delta_4$ and δ_5 are all positive, as expected (see Section II). Second, all coefficients are significant at the 5 per cent level. Third, the long-run effect of the capital stock on the unemployment rate is not significantly different from zero (*p*-value = 0.13). This finding seems to agree with most of the literature (see, e.g., Blanchard and Summers, 1986, p. 27 and Layard, Nickell and Jackman, 1991, p. 107). The short-run effect is negative and significant, however: the coefficient of k_{it} is −41.58 (t = −8.08, *p*-value = 0). Since k_{it} is measured in logs, this coefficient implies that a *ceteris paribus* increase in the current period's capital stock by 1 per cent is expected to cause a decrease in current period's unemployment rate by about 0.42 of a percentage point (see Wooldridge, 2006, pp. 720-721).

Table 2: *Fixed-effects Estimation of Equation (1) without a Time Trend*

Part A. The 19-country panel
Regression Coefficients and their t-Statistics

Constant	k_{it}	$k_{i,t-1}$	gun_{it}	f_{it}	gsb_{it}	gnm_{it}	$u_{i,t-1}$	$u_{i,t-2}$
0.55***	−41.58***	42.06***	0.02**	0.17**	0.08***	−0.000008***	1.28***	−0.37***
(7.98)	(−8.08)	(8.26)	(2.25)	(2.46)	(9.63)	(−3.58)	(18.63)	(−4.59)

Goodness of Fit and Tests of Hypotheses

$\bar{R}^2 = 0.93$, *RESET*: $\chi_1^2 = 0.04$ [0.84], $\chi_2^2 = 1.15$ [0.56]

H_0: Consistency of the fixed-effects estimators (tested by the Durbin-Wu-Hausman test)

k_{it}: *t*-stat = −0.92 [0.36]; gun_{it}: *t*-stat = 0.90 [0.37]; gsb_{it}: *t*-stat = 1.40 [0.16]; gnm_{it}: *t*-stat = −0.95 [0.34]

H_0: $\delta_1 + \delta_2 = 0$ (or $\eta = 0$), $\chi_1^2 = 2.29$ [0.13]

Part B. The 14-Country Panel
Regression Coefficients and their t-Statistics

Constant	k_{it}	$k_{i,t-1}$	gun_{it}	f_{it}	gsb_{it}	gnm_{it}	$u_{i,t-1}$	$u_{i,t-2}$
0.55***	−17.86***	18.09***	0.02*	0.15**	0.07***	−0.000008***	1.31***	−0.40***
(7.16)	(−7.80)	(8.00)	(1.72)	(2.27)	(9.87)	(−3.71)	(16.57)	(−4.33)

Goodness of Fit and Tests of Hypotheses

$\bar{R}^2 = 0.93$, *RESET*: $\chi_1^2 = 0.06$ [0.81], $\chi_2^2 = 1.95$ [0.38]

H_0: Consistency of the fixed-effects estimators (tested by the Durbin-Wu-Hausman test)

k_{it}: *t*-stat = −1.32 [0.19]; gun_{it}: *t*-stat = 1.19 [0.23]; gsb_{it}: *t*-stat = 1.75* [0.08]; gnm_{it}: *t*-stat = −1.27 [0.20]

H_0: $\delta_1 + \delta_2 = 0$ (or $\eta = 0$), $\chi_1^2 = 2.01$ [0.16]

Notes: (1) ***, **, * denote statistical significance at the 1 per cent, 5 per cent and 10 per cent level respectively; (2) the numbers in parentheses underneath the regression coefficients are *t*-statistics, whereas those in square brackets following the values of the test statistics are *p*-values; (3) the hypothesis of consistency of our fixed-effects estimators is tested using the Durbin-Wu-Hausman test, by treating as endogenous the variable that precedes the test statistic (*t*-stat); (4) the hypothesis H_0: $\delta_1 + \delta_2 = 0$ implies that the long-run effect of the capital stock on the observed unemployment rate is zero; (5) the results of this table have been produced by the econometric program *RATS* v. 7 with the option that produces "clustered" standard errors, which are robust to heteroscedasticity and to autocorrelation (see *RATS Reference Manual*, p. 298).

Fourth, a *ceteris paribus* increase in the growth rate of the percentage of unionised workers by one percentage point causes an increase in the

unemployment rate by 0.02 of a percentage point. Fifth, a *ceteris paribus* increase in the index of firing restrictions by one unit causes an increase in the unemployment rate by 0.17 of a percentage point. This estimate can be compared with those found by HK (2010), which range from 1.18 to 1.93; the latter are much larger, since, as we noted earlier, they refer to the effect of this index on the NAIRU and not just on the current period's unemployment rate.

Sixth, a *ceteris paribus* increase in the growth rate of social-security-benefits by one percentage point causes an increase in the unemployment rate by 0.08 of a percentage point. Again, compare this estimate with those found by HK (2010), 0.15 and 0.26. Seventh, a *ceteris paribus* increase in the growth rate of net immigration by one percentage point causes a decrease in the unemployment rate by 0.000008 of a percentage point. Once again, compare the coefficient –0.000008, which is statistically significant ($t = -3.58$, *p*-value = 0), but economically *in*significant, with those obtained by HK (2010), namely –0.0002 and –0.0019. Eighth, the coefficients of the lagged values of u_{it} are large and highly significant, reflecting the strong "lagged adjustment processes" that are at work in the labour market (see Karanassou, Sala and Salvador, 2008, p. 984).

Now, consider the results of our diagnostic tests. First, the value of $\bar{R}^2$ is 0.93, which suggests a good fit. Second, the equation passes easily the *RESET*, which is implemented by adding to the set of explanatory variables the squared fitted values of u_{it} (*p*-value = 0.84) as well as both the squared and the cubed fitted values (*p*-value = 0.56). Thus, there is no evidence against the specification of Equation (1).

Third, Equation (1) also passes easily the Durbin-Wu-Hausman (DWH) test of consistency of the standard fixed-effects estimator (see Davidson and MacKinnon, 1993, pp. 237-242). To implement this test, we first choose a set of instrumental variables (IVs), which contains only lagged values of the variables that appear in Equation (1), namely $\{1, k_{i,t-1}, gun_{i,t-1}, f_{i,t-1}, gsb_{i,t-1}, gnm_{i,t-1}, t, u_{i,t-1}, u_{i,t-2}\}$. Next, we run fixed-effects regressions, each of which has these IVs as explanatory variables and one of the following variables as the dependent variable: k_{it}, gun_{it}, gsb_{it}, or gnm_{it}. The series of the fitted values from each of these regressions is then included (one at a time) as an additional explanatory variable in Equation (1) and the *t*-statistic of its coefficient is calculated. The test does not reject at any conventional level, as the *p*-values of these *t*-statistics range from 0.16 to 0.37 (see Part A of Table 2). The reason why we carry out the DWH test is that the above four variables (k_{it}, gun_{it}, gsb_{it} and gnm_{it}) could potentially be endogenous, i.e., they could be correlated with the error term. For example, an adverse shock to the unemployment rate (i.e., an increase in ε_{it}) may cause an increase in the rate of growth of social-security benefits

(i.e., an increase in gsb_{it}); it may also cause a decrease in the influx of immigrants (i.e., a decrease in gnm_{it}).[3]

Fourth, the 14-country panel gives similar results, the only noticeable difference being in the size of the coefficients of k_{it} and k_{it-1} (see Part B of Table 2). Thus, there is no strong evidence of sample-selection bias or major lack of robustness to substantial changes in the sample. All in all, the diagnostic tests enhance the reliability of our results.

V CONCLUSIONS

Our goal in this paper has been to obtain reliable estimates of the effects on the unemployment rate of certain variables, such as the capital stock, wage-push factors and net immigration. To achieve this goal, we estimate a reduced-form equation for the unemployment rate derived from a dynamic labour-market model that incorporates its principal characteristics.

Using a dynamic fixed-effects model for a panel of nineteen OECD countries for the time period 1970-2008, we find, first, that wage-push factors, such as increases in union power and in employment protection raise the unemployment rate. Thus, the governments of the OECD countries considered here could reduce the unemployment rate by easing employment protection laws and other rigidities (see OECD, 1994). Second, the current unemployment rate is strongly and positively influenced by its value of the previous period. Thus, the governments of the countries in the sample have an additional incentive to fight current unemployment (using standard policy measures), to prevent it from becoming prolonged. Third, the long-run effect of the capital stock on the unemployment rate is zero. Fourth, legal immigration reduces the unemployment rate, but this effect is practically negligible. To escape possible criticisms of sample-selection bias or lack of robustness to substantial changes in the sample, we also estimate our equation after removing from the panel five countries in which the index of firing restrictions exhibits little or no variability, but obtain similar results.

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[3] Although the DWH test suggests that none of the above four explanatory variables (k_{it}, gun_{it}, gsb_{it}, gnm_{it}) can be considered endogenous, we have also estimated Equation (1) by the IV method, using the above set of IVs augmented with the current values of k_{it}, gun_{it}, f_{it}, gsb_{it}, and gnm_{it}, since the DWH test suggests that these variables may be considered exogenous. The IV estimates are remarkably similar to those of least squares, so we do not report them. The only noticeable difference is that the estimated coefficient of gun_{it}, which is again 0.02, is significant only at the 10 per cent level, since its p-value is 0.07. Note that the IV regression easily passes the standard over-identifying restrictions test (p-value = 0.21).

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