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IS THE IMPACT OF LABOR TAXES ON UNEMPLOYMENT ASYMMETRIC?

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This paper tests whether increases and decreases in labor taxes have an asymmetric impact on unemployment. Using a panel of 16 OECD countries over the period 1970–2005, we estimate a panel unobserved-component model to account for the fact that unemployment rates and labor taxes are nonstationary but not cointegrated. We find a positive impact of labor tax increases on unemployment in European and Nordic countries, whereas for labor tax decreases, no significant impact is found in these countries. For Anglo-Saxon countries, neither increases nor decreases in labor taxes have any impact on unemployment.

Keywords: Unemployment, Labor Taxes, Asymmetry, Unobserved-Component Model

1. INTRODUCTION

High and persistent unemployment in many OECD countries is one of the biggest challenges for policymakers and labor economists in recent times. It is a widespread belief, especially among policymakers, that the increase in labor taxes over the past decades is one of the prime factors responsible for the increase in unemployment. Consequently, the alleviation of the high tax burden on labor has been declared to be one of the prime instruments for fighting high unemployment. This strategy relies on the assumption that the alleged impact of labor taxes on unemployment is more or less symmetric; i.e., cutting taxes will reduce unemployment as much as their increase induced unemployment to go up.

In standard labor market models, the impact of an increase in labor taxes on unemployment depends on the degree to which employees succeed in shifting the higher tax burden onto the employer. This shifting forward is only possible if alternative income sources, e.g., unemployment benefits, are not equally affected

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by the increase in taxes [Pissarides (1998); Nickell and Layard (1999); Daveri and Tabellin (2000)]. In this case, labor taxes have a positive impact on unemployment, as they drive a wedge between labor income and alternative income. The extent of this impact crucially depends on the amount of labor market competition. Excessive labor market regulations (e.g., extensive employment protection and high minimum wages), high union bargaining power, and insider behavior of employed workers all obstruct competition on the labor market, implying that a higher proportion of taxes are shifted forward to labor costs. Calmfors and Driffill (1988) have argued, though, that both in highly centralized/coordinated wage bargaining systems and in fully decentralized/competitive systems, unions are likely to take a more moderate stand in response to adverse shocks, e.g., tax increases, hitting the economy.

Surveying the empirical literature, the estimated elasticity of unemployment with respect to labor taxes ranges from zero [Bean et al. (1986); Nickell (1997); Blanchard and Wolfers (2000); Layard et al. (2005)] over medium-sized [Elmeskov et al. (1998); Nickell and Layard (1999); Nickell et al. (2005); Planas et al. (2007); Berger and Everaert (2010)] up to large [Daveri and Tabellin (2000)]. All these studies assume the impact of taxes to be symmetric. The contribution of this paper is to test whether the impact of labor taxes on unemployment is asymmetric. To the best of our knowledge, this has not yet been tested empirically.

The plan of the paper is as follows: Section 2 gives reasons for why the labor taxes–unemployment trade-off might be asymmetric. Section 3 lays out the empirical model and presents the results. Section 4 concludes.

2. ASYMMETRIES IN THE LABOR TAXES–UNEMPLOYMENT TRADE-OFF

A first theoretical justification for asymmetries in the labor taxes–unemployment trade-off is the insider–outsider distinction [see, e.g., Blanchard and Summers (1986); Lindbeck and Snower (1987)]. It states that insiders (those currently employed) are insulated from competition by outsiders (those currently unemployed) due to labor turnover costs (i.e., costs associated with hiring, training, and firing). As a result, insiders have scope to push wages above the market-clearing level; i.e., they may use their privileged position to shift the tax burden to the employer (resulting in higher unemployment) in response to an increase in taxes and try to push for higher net wages (instead of lower unemployment) in response to a decrease in taxes.[1] It is often argued, though, that insider–outsider effects are less strong when unions are centralized (as they are in, e.g., the Nordic countries). There are at least two reasons for this. First, higher wages in one sector of the economy cause unemployment to rise and consequently imply (i) a fall in output and in the taxable base and (ii) higher costs for unemployment benefits. These negative wage externalities will be internalized in a centralized wage-setting system [Calmfors and Driffill (1988)],[2] giving rise to more moderate wage claims. Second, membership rules in centralized wage settings are more favorable to the unemployed, who typically remain union members. Thus they receive more

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attention in the wage bargaining process, implying that insiders have less incentive to push wages above the market-clearing level [see, e.g., Blanchard and Summers (1986), Layard et al. (2005)].[3] Insider–outsider effects in wage formation may therefore be a less appropriate explanation for potential asymmetries in the labor taxes–unemployment trade-off in countries with a centralized wage-bargaining system.

A second theoretical justification for asymmetries in the labor taxes–unemployment trade-off is the occurrence of asymmetric adjustment costs. Different costs for hiring and firing imply different speeds of adjustment for employment and hence unemployment in response to a labor market shock. A large literature provides considerably empirical evidence in support of asymmetries in labor demand. Using various techniques and functional forms, Burgess (1992a, 1992b), Pfann and Palm (1993), Hamermesh and Pfann (1996), and Holly and Turner (2001) show that there are asymmetries in labor demand due to asymmetric adjustment costs.

3. EMPIRICAL ANALYSIS

3.1. Data

Our data set consists of yearly observations for 16 OECD countries over the period 1970–2005. The unemployment rate is taken from the OECD Economic Outlook. As a measure of labor taxes, we use an update of the effective tax rates on employed labor from Martinez-Mongay (2003).[4] This tax rate has been calculated with the so-called Mendoza–Razin–Tezar approach [see Mendoza et al. (1994)] using the EU AMECO database. Figure 1 shows that labor taxes steadily increased in most OECD countries until the mid-1990s. Only recently, a decrease in labor taxes can be observed in a number of countries, most notably in Finland, Ireland, and the Netherlands. With the availability of these recent data, it is now possible to analyze the response of unemployment to decreasing labor taxes.

3.2. Empirical Specification: An Unobserved-Component Approach

Empirical studies on the determinants of unemployment typically estimate a reduced-form unemployment equation linking the rate of unemployment to various labor market institutions and macroeconomic shocks [see, e.g., Nickell et al. (2005) for a recent example]. One major concern with this approach is that observed unemployment rates are found to exhibit unit root behavior in most OECD countries over the past four decades.[5] Thus, unless there is a cointegrating relation between unemployment and its alleged determinants, standard estimation methods yield spurious results. Using a panel of yearly data for 16 OECD countries ranging from 1960 to 1995, Berger and Everaert (2009) show that unemployment is not cointegrated with a large set of labor market institutions and macroeconomic shocks. The finding of no panel cointegration does not imply that there is no relation between unemployment and labor market institutions, though. Economic

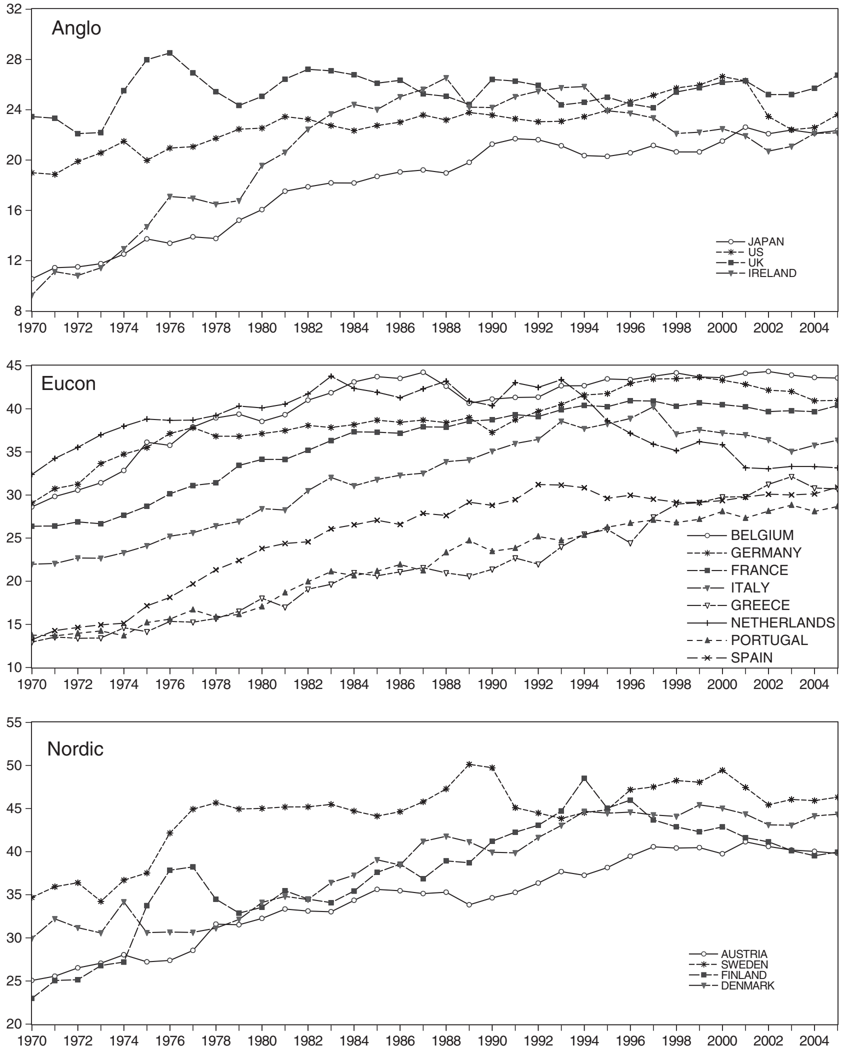

FIGURE 1. Labor taxes.

theory relates the equilibrium rate of unemployment to a wide variety of factors, some difficult to measure or even unobservable, e.g., the reservation wage, which is a function of, among others, the value of leisure. By inducing a unit root component in the residuals, both missing nonstationary variables and measurement errors in nonstationary variables turn an otherwise cointegrating relation into a spurious regression [see Everaert (2007) for a simulation experiment]. To solve this missing-variables problem, Planas et al. (2007) and Berger and Everaert (2010) set up an unobserved-component model in which the sum of all missing variables is treated as a latent state variable and identified through the Kalman filter. Here, we follow Berger and Everaert, but extend their model to allow for an asymmetric impact of labor taxes on unemployment.

Let the total unemployment rate u_{it} be the sum of an equilibrium component u^*_{it}, which is a function of structural factors driving long-run unemployment, and a temporary component u^c_{it}, which we label cyclical unemployment:

$$u_{it} = u^*_{it} + u^c_{it}, \qquad i = 1, \ldots, N, \qquad t = 1, \ldots, T, \tag{1}$$

where N is the number of countries and T is the number of time series observations. To allow cyclical unemployment to exhibit the standard hump-shaped pattern, u^c_{it} is assumed to be an AR(2) process:

$$u^c_{it} = \phi_1 u^c_{i,t-1} + \phi_2 u^c_{i,t-2} + \eta^c_{i,t-1}, \qquad \eta^c_{i,t-1} \sim NID\big(0, \sigma^2_{\eta^c_i}\big). \tag{2}$$

The equilibrium rate u^*_{it} is assumed to be given by

$$u^*_{it} = u^*_{it-1} + \varepsilon_{it}, \tag{3}$$

such that ε_{it} reflects all factors that induce a permanent shift in the equilibrium rate of unemployment. To estimate the impact of labor taxes on u^*_{it} we disentangle ε_{it} into the impact of labor taxes ε^T_{it} and the impact of all other factors ε^Z_{it}; i.e., $\varepsilon_{it} = \varepsilon^T_{it} + \varepsilon^Z_{it}$, where

$$\varepsilon^T_{it} = \beta_1 \Delta\mathrm{TAX}_{it} + \beta_2 D_{it} \Delta\mathrm{TAX}_{it}, \tag{4}$$

$$\varepsilon^Z_{it} = \delta \varepsilon^Z_{it-1} + \eta^Z_{i,t-1}, \qquad \eta^Z_{i,t-1} \sim \mathrm{NID}\big(0, \sigma^2_{\eta^Z_i}\big), \tag{5}$$

with $D_{it} = 1$ if $\Delta\mathrm{TAX}_{it} < 0$ and zero otherwise. Thus, the impact of labor taxes on u_{it} is measured by β_1 and β_2 in (4). Accounting for a potential asymmetric impact, β_1 measures the impact of increasing labor taxes, whereas β_2 measures the differential impact of decreasing labor taxes on unemployment. Thus, $\beta_1 + \beta_2$ measures the effect of decreasing labor taxes on unemployment. If there are asymmetric effects, one would expect $\beta_2 \neq 0$, in particular $\beta_2 < 0$. Equation (5) describes the stochastic process of the other determinants of equilibrium unemployment. As a pure random walk process would result in a nonsmooth series that is hard to reconcile with the expected smooth evolution of the structural characteristics driving equilibrium unemployment, the AR(1) specification in (5) allows a smooth evolution of ε^Z_{it} over time; i.e., the closer δ to one, the smoother ε^Z_{it}. If $\delta = 0$, ε^Z_{it} is a pure random walk process. Note that in order to induce smoothness, the equilibrium rate of unemployment is nowadays often modeled as an I(2) series; i.e., δ is set to one [see, e.g., Orlandi and Pichelmann (2000)]. We do not restrict δ to be equal to one in (5) as in this case ε^Z_{it}, and therefore also u_{it}, would exhibit a (time-varying) drift, which is hard to justify from an economic perspective.

The model in (1)–(5) can be written in a panel linear Gaussian state space representation, from which the unobserved states can be identified using the Kalman filter and the unknown parameters can be estimated using maximum likelihood (see Appendix A for more details).

3.3. Country Grouping

In order to take into account possible heterogeneity of the impact of taxes on unemployment, we group countries according to their wage-setting institutions. Following Daveri and Tabellin (2000), Doménech and García (2008), and Berger and Everaert (2010), and using the same notation, we classify countries into three different groups. The empirical specification, outlined in (1)–(5), is estimated separately for each of these three groups; i.e., the parameters ϕ_1, ϕ_2, β_1, β_2 and δ are assumed to be homogeneous within each of the three country groups considered, but heterogeneous over these groups. The first group (NORDIC) includes Austria, Denmark, Finland, and Sweden. These countries are characterized by strong unions, wage bargaining at a central level, and/or a high degree of coordination. The unemployment incidence of labor taxes is expected to be moderate in these countries. The second group (EUCON) includes Belgium, France, Germany, Italy, the Netherlands, Portugal, Spain, and Greece. In these countries, wages are generally bargained at the intermediate level, without a strong tendency to coordination across bargaining units. In this setting, unions are expected to use their bargaining power to shift the burden of higher labor taxes onto employers. The third group (ANGLO) includes Japan, Ireland, the United States, and the United Kingdom. In these countries unions are typically not strong enough to shift the tax burden.

3.4. Results

Before looking at the specific coefficient estimates, we perform some diagnostic tests. In the unobserved-component model presented in (1)–(5), the innovations $\eta^c_{i,t-1}$ and $\eta^Z_{i,t-1}$ are assumed to be white Gaussian noise. Following Durbin and Koopman (2001) we check whether this property holds by testing for autocorrelation and nonnormality in the standardized one–step ahead prediction errors v_{it} of the state space model. At the country group level, we use two LM tests for autocorrelation suggested by Baltagi and Li (1995). The first test specifies the residuals as an AR(1) process, i.e., $v_{it} = \rho v_{it-1} + \varepsilon_{it}$, and tests the null hypothesis that $\rho = 0$. The second test models the residuals as an MA(1) process, i.e., $v_{it} = \varepsilon_{it} + \lambda\varepsilon_{it-1}$, and tests the null hypothesis that $\lambda = 0$. The results are reported in Table 1. Both tests show that we cannot reject the null of no autocorrelation in any of the three country groups. More detailed country-specific diagnostic tests, including normality tests and higher-order autocorrelation tests using Ljung–Box Q-tests, are reported in Appendix B. For the majority of countries the null hypotheses of normality and no autocorrelation cannot be rejected.

The parameter estimates are presented in Table 1. For the ANGLO group we do not find a clear significant impact of labor taxes on unemployment either for an increase or for a decrease in taxes. In contrast, there is a significant positive

TABLE 1. Parameter estimates

	ANGLO $N = 4$	EUCON $N = 8$	NORDIC $N = 4$
β_1	0.01	0.16***	0.07**
	(0.07)	(0.04)	(0.03)
β_2	−0.23	−0.15**	−0.07
	(0.15)	(0.08)	(0.07)
ϕ_1	0.87***	1.48***	1.28***
	(0.24)	(0.06)	(0.21)
ϕ_2	−0.39	−0.76***	−0.71***
	(0.25)	(0.06)	(0.15)
δ_1	0.62***	0.54***	0.54***
	(0.16)	(0.07)	(0.18)
$t_{\beta_2<0}$	−1.53*	−1.87**	−0.99
	[0.06]	[0.03]	[0.16]
$t_{\beta_1+\beta_2\neq 0}$	−0.95	0.02	−0.04
	[0.34]	[0.98]	[0.97]
$\mathrm{LM}_{\mathrm{AR}}$	0.10	0.79	0.08
	[0.75]	[0.07]	[0.78]
$\mathrm{LM}_{\mathrm{MA}}$	0.31	0.27	0.28
	[0.62]	[0.61]	[0.61]

Note: $t_{\beta_2<0}$ is a one-sided t-test for the null hypothesis $H_0 : \beta_2 = 0$ against the alternative $H_1 : \beta_2 < 0$. $t_{\beta_1+\beta_2\neq 0}$ is a two-sided t-test for the null hypothesis $H_0 : \beta_1 + \beta_2 = 0$ against the alternative $H_1 : \beta_1 + \beta_2 \neq 0$. $\mathrm{LM}_{\mathrm{AR}}$ and $\mathrm{LM}_{\mathrm{MA}}$ are LM tests for an AR(1) respectively MA(1) structure in the one-step ahead prediction errors (null hypothesis is no autocorrelation). Standard errors are in parentheses. p-values are in brackets.
*, ** and *** Significance at the 10%, 5%, and 1% level, respectively (using a two-sided t-test).

impact of increases in labor taxes on unemployment in both the EUCON and the NORDIC group, with the impact being more moderate in the latter. Moreover, in the EUCON group there is clear evidence of an asymmetric impact of labor taxes; i.e., the differential impact of tax decreases β_2 is significantly negative (as indicated by the one-sided test $t_{\beta_2<0}$), with the total impact of tax decreases $\beta_1+\beta_2$ not being significantly different from zero (as indicated by the two-sided test $t_{\beta_1+\beta_2\neq 0}$). In the NORDIC group, there is less evidence of asymmetry, as the point estimate of β_2 is negative but not statistically significantly different from zero. However, as in the other two groups, the impact of tax decreases is also not statistically significantly different from zero. As far as the other coefficients are concerned, the estimates of δ_1 indicate that equilibrium unemployment is considerably smoother than a simple random walk but not as smooth as an $I(2)$ process. Further, ϕ_1 and ϕ_2 imply that cyclical unemployment exhibits the standard hump-shaped pattern in all country groups.

FIGURE 2. Labor taxes and their contribution to unemployment (as measured by cumulative ε_{it}^{T}) for the Netherlands.

Figure 2 visualizes the asymmetry for the Netherlands. It shows the evolution of labor taxes and their contribution to unemployment, as measured by the cumulated values of ε_{it}^{T}. We have chosen to graph data for the Netherlands because this country (i) shows a notable decline in labor taxes since 1994 and (ii) belongs to the EUCON group, where there is clear evidence of asymmetry. Labor taxes in the Netherlands increased from about 32.5% in 1970 to around 43% in 1993. This induced a rise in unemployment with 2.7% points. Due to the asymmetric impact, the subsequent strong decline in labor taxes to 33% by 2005 did not lead unemployment to go down, though. For the average of the country group EUCON (graph not included), the increase in labor taxes from 22% to 36% between 1970 and 1998 accounts for a 2.25% increase in the average unemployment rate. Again, the slight decline in labor taxes after 1998 did not lead unemployment to go down.

4. CONCLUSION

This paper tests whether the impact of labor taxes on unemployment is asymmetric. The insider–outsider hypothesis and the occurrence of asymmetric adjustment costs both give reasons to suspect that the impact on unemployment is stronger for increases than for decreases in labor taxes. We estimate a panel unobserved-component model to account for the fact that unemployment rates and labor taxes are nonstationary but not cointegrated, due to unobserved nonstationary variables driving unemployment. The sixteen OECD considered countries are divided into three groups to allow for heterogeneity in the labor taxes–unemployment trade-off depending on their labor market characteristics. The results show that labor taxes do not affect unemployment in the Anglo-Saxon countries. Increasing labor taxes have a significant positive impact on unemployment in the European and Nordic countries, with the impact being more moderate in the latter. In both these country groups, the impact of tax decreases is smaller than that of increases and is not significantly different from zero. These results imply that the strategy of reducing labor taxes to fight high unemployment is not too promising.

NOTES

1. Huizinga and Schiantarelli (1992) study the insider–outsider distinction within a general equilibrium model and confirm that it gives rise to an asymmetric response of employment in expansions and recessions.

2. In addition to this fiscal externality, there are at least six other negative wage externalities. See Calmfors (1993) for an overview.

3. However, as pointed out by Calmfors (1993), only employed union members elect the union officials who represent the union in the wage bargaining process. The advantage of being a union member for the unemployed is therefore not obvious.

4. We would like to thank Carloz Martinez-Mongay for providing this data set.

5. Because the unemployment rate is bounded between zero and unity, one would expect it to be stationary. Over longer periods of time, this indeed seems to be the case. Postwar data, though, shows a clear upward trend and strong persistence, resulting in nonstationary behavior as typically documented by unit root tests. This small-sample behavior requires the unemployment rate to be treated as a nonstationary process. Note that this is in line with the current practice in the literature [see, e.g., Apel and Jansson (1999a, 1999b); Fabiani and Mestre (2004)].

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APPENDIX A: STATE SPACE REPRESENTATION

The model in (1)–(5) can be written in a panel linear Gaussian state space representation, which consists of an observation and a state equation. The observation equation models the vector of observed unemployment rates $u_t = [u_{1t}, \ldots, u_{Nt}]'$ as a function of a vector of unobserved states α_t. More specifically,

$$\begin{bmatrix} u_{1t} \\ \vdots \\ u_{Nt} \end{bmatrix} = \begin{bmatrix} I_N & O_N & I_N & O_N & O_N & O_N \end{bmatrix} \begin{bmatrix} u_t^c \\ u_{t-1}^c \\ u_t^* \\ \varepsilon_t^Z \\ \varepsilon_t^T \\ \iota_T \end{bmatrix}, \qquad \textbf{(A.1)}$$

$$u_t \quad = \quad Z \quad \alpha_t,$$

where I_N is the identity matrix of order N, O_N is an $N \times N$ matrix of zeros, and ι_T is an $N \times 1$ column vector of ones. The unobserved states in the state vectors $u_t^c = [u_{1t}^c, \ldots, u_{Nt}^c]'$, $u_t^* = [u_{1t}^*, \ldots, u_{Nt}^*]'$, $\varepsilon_t^T = [\varepsilon_{1t}^T, \ldots, \varepsilon_{Nt}^T]'$, and $\varepsilon_t^Z = [\varepsilon_{1t}^Z, \ldots, \varepsilon_{Nt}^Z]'$ are defined in (2), (3), (4), and (5) respectively. The unobserved states are modeled in the state equation

$$\begin{bmatrix} u_{t+1}^c \\ u_t^c \\ u_{t+1}^* \\ \varepsilon_{t+1}^Z \\ \varepsilon_{t+1}^T \\ \iota_T \end{bmatrix} = \begin{bmatrix} \phi_1 I_N & \phi_1 I_N & O_N & O_N & O_N & O_N \\ I_N & O_N & O_N & O_N & O_N & O_N \\ O_N & O_N & I_N & I_N & I_N & O_N \\ O_N & O_N & O_N & \delta I_N & O_N & O_N \\ O_N & O_N & O_N & O_N & O_N & \beta_t \\ O_N & O_N & O_N & O_N & O_N & I_N \end{bmatrix} \begin{bmatrix} u_t^c \\ u_{t-1}^c \\ u_t^* \\ \varepsilon_t^Z \\ \varepsilon_t^T \\ \iota_T \end{bmatrix} + \begin{bmatrix} I_N & O_N \\ O_N & O_N \\ O_N & O_N \\ O_N & I_N \\ O_N & O_N \\ O_N & O_N \end{bmatrix} \begin{bmatrix} \eta_t^c \\ \eta_t^Z \end{bmatrix}, \qquad \textbf{(A.2)}$$

$$\alpha_{t+1} \quad = \quad S \quad \alpha_t \quad + \quad R \quad \eta_t,$$

where $\beta_t = \beta_1 \mathrm{diag}\,[\Delta Tax_{1t} \ldots \Delta Tax_{Nt}] + \beta_2 \mathrm{diag}\,[D_{1t}\Delta Tax_{1t} \ldots D_{Nt}\Delta Tax_{Nt}]$, $\eta_t^c = [\eta_{1t}^c, \ldots, \eta_{Nt}^c]'$, and $\eta_t^Z = [\eta_{1t}^Z, \ldots, \eta_{Nt}^Z]'$. We assume that the innovations in η_t are mutually independently normally distributed with variances that are heterogeneous over countries; i.e., $\eta_t \sim N(0, Q)$, with

$$Q = \begin{bmatrix} \mathrm{diag}\big(\sigma_{\eta^c}^2\big) & O_N \\ O_N & \mathrm{diag}\big(\sigma_{\eta^Z}^2\big) \end{bmatrix}, \qquad \textbf{(A.3)}$$

where $\sigma_{\eta^c}^2 = [\sigma_{\eta_1^c}^2, \ldots, \sigma_{\eta_N^c}^2]'$ and $\sigma_{\eta^Z}^2 = [\sigma_{\eta_1^Z}^2, \ldots, \sigma_{\eta_N^Z}^2]'$.

For given values of the unknown parameters, the unobserved states α_t and the log likelihood of the model in equations (A.1) and (A.2) can be calculated by a routine application of the Kalman filter from which the unknown parameters can be estimated using maximum likelihood [see, e.g., Harvey (1989); Durbin and Koopman (2001)]. The stationary state variables $(u_{it}^c, u_{i,t-1}^c, \varepsilon_{it}^Z, \varepsilon_{it}^T)$ are initialized by drawing from their stationary distributions, whereas a diffuse initialization is used for the nonstationary state variables (u_{it}^*). Standard errors for the parameter estimates are calculated by inverting the Hessian matrix.

APPENDIX B: COUNTRY-SPECIFIC DIAGNOSTIC TESTS

	JB	$Q(1)$	$Q(2)$	$Q(3)$	$Q(4)$
Ireland	0.35 [0.84]	0.00 [0.99]	3.70 [0.16]	6.56 [0.09]	6.62 [0.16]
Japan	3.11 [0.21]	4.84 [0.03]	12.98 [0.00]	16.57 [0.00]	20.67 [0.00]
U.K.	1.60 [0.45]	6.15 [0.01]	7.13 [0.03]	8.89 [0.03]	10.25 [0.04]
U.S.	12.63 [0.00]	0.00 [0.99]	0.30 [0.86]	0.55 [0.91]	0.66 [0.96]
Belgium	0.22 [0.90]	1.27 [0.26]	2.85 [0.24]	2.89 [0.41]	2.99 [0.56]
France	2.40 [0.30]	0.17 [0.68]	1.35 [0.51]	1.63 [0.65]	1.83 [0.77]
Germany	0.61 [0.74]	0.18 [0.68]	2.62 [0.27]	2.66 [0.45]	2.76 [0.60]
Greece	3.75 [0.15]	1.75 [0.19]	1.81 [0.40]	2.07 [0.56]	2.08 [0.72]
Italy	0.65 [0.72]	0.59 [0.44]	0.94 [0.63]	0.98 [0.81]	1.74 [0.78]
Netherlands	0.99 [0.61]	3.72 [0.05]	4.07 [0.13]	5.06 [0.17]	6.19 [0.19]
Portugal	0.11 [0.94]	0.12 [0.72]	0.13 [0.93]	2.68 [0.44]	2.75 [0.60]
Spain	1.58 [0.45]	1.72 [0.19]	1.81 [0.40]	1.87 [0.60]	1.88 [0.76]
Austria	1.04 [0.60]	0.07 [0.79]	5.11 [0.08]	5.12 [0.16]	5.32 [0.26]
Finland	0.42 [0.81]	0.10 [0.75]	1.84 [0.40]	3.36 [0.34]	4.24 [0.38]
Denmark	5.92 [0.05]	4.46 [0.04]	5.74 [0.06]	9.20 [0.03]	9.46 [0.05]
Sweden	2.72 [0.26]	0.00 [0.99]	1.08 [0.58]	1.45 [0.69]	3.24 [0.52]

Note: *JB* is the Jarque–Bera test for normality of the one–step ahead prediction errors (null hypothesis is normality). $Q(m)$ is the Ljung–Box portmanteau test for autocorrelation at lag m (null hypothesis is no autocorrelation). p-values are in brackets.