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# Employment adjustment in two countries with poor reputations: Analysis of aggregate, firm, and flow data for Portugal and Germany

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**Abstract.** This paper supplements aggregate time-series analysis of the speed of employment adjustment with evidence from firm panel and flow data for two countries – Portugal and Germany – sharing unenviable labor market reputations. The Portuguese labor market is often portrayed as terminally inert, while that of Germany as badly ailing. We report broad consistency in the results across data sets in favor of Portugal. In benchmarking Portugal against Germany, the adverse reputation of the former – if not necessarily that of the latter country – may have been exaggerated in contemporary policy debate.

**Key words:** employment adjustment, employment protection, error correction model, panel estimation, labor reallocation

**JEL Classification:** C22, C23, J23

## 1 Introduction

Certain practicalities have long ensured academic interest in the process of employment adjustment in the labor market. Thus, the presence of firm-specific training introduces employment adjustment costs, ensuring that firms will not expand/contract employment in immediate response to increased/decreased sales. Other elements of fixity produced by the costs of the personnel function and government-mandated benefits such as health and pension programs yield the same result. More recently, with the erection of elaborate systems of employment protection in many countries, attention has shifted away from hiring costs to firing costs. This concern led to liberalizing moves in some European countries in the 1980s, but the partial nature of such reforms and the stubborn persistence of unemployment has ensured

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that the level of firing costs continues to preoccupy policy debate and to motivate most of the employment adjustment research.

Considerable research effort has, then, been devoted to determining the effect of employment protection on levels of employment (and unemployment) and on the speed of adjustment of labor demand to changes in output (see, *inter al.*, Abraham and Houseman 1994, Addison *et al.* 2000, Bentolila and Bertola 1990, Blanchard and Portugal 2001, Botero *et al.* 2003, Burgess 1988, Garibaldi 1998, Hamermesh 1993b, OECD 1999, Scarpetta 1996). Theory suggests that employment protection should lower the speed of adjustment of employment to changes in output, even if the effects on levels of employment and unemployment are muddied by the opposing effects on labor flows created by a policy of employment protection (Addison and Teixeira 2003a, b).

In the present treatment, we focus on the process of labor market adjustment in two countries: Portugal and Germany. For its part, Portugal provides an interesting case study because of its reputation as an exemplar of over-ambitious employment protection. The country has been consistently ranked by the OECD as one of the most rigid labor markets (e.g. OECD 1994, Table 6.7, Panel B; OECD 1999, Table 2.6). But we seek less to contest that reputation head-on than to build up a picture of the nature of employment adjustment by filtering information from separate types of data: aggregate, firm level, and flow data. In checking for consistency in the evidence across tiers of data, we will benchmark Portugal against Germany. Germany possesses a similar (medium to high) degree of centralization and degree of co-ordination of bargaining (see Elmeskov *et al.* 1998, Table B2). Despite its ailing labor market, Germany has nevertheless enjoyed a consistently more favorable regulatory reputation than Portugal.

We begin this empirical inquiry by considering aggregate time-series data, traditionally used in the analysis of employment determination over time. It is standard in such treatments to apply an error correction model to investigate the dynamic properties of labor demand. As the first step in our own inquiry, therefore, we report results from a (single-stage) error correction model to simulate the employment adjustment path implied by an exogenous shock. But because time-series evidence may hide what is going on at the level of the firm – for example, if labor demand adjustment at the micro level is lumpy, and the process unsynchronized, the resulting aggregate behavior may nevertheless appear to be quite smooth – there is an obvious case for supplementing this type of evidence with information gathered at a lower level of aggregation.

To this end, we will also assemble two micro data sets: a firm-level, manufacturing data set for Portugal based on balance sheet records, and an establishment-level, economy-wide data set for Germany based on an employment survey. We will use these data not merely to derive a set of panel estimates that may be compared with our time-series evidence but also to provide information on job and worker flows. Two pieces of information from our panels will prove useful in this regard. First, we will use employment changes at the firm level to measure job creation/destruction as well as the corresponding transition probabilities between job creation and job destruction regimes. Second, we will use flows of newly employed and separated workers to compute hiring and separation rates. We anticipate that a smooth process of adjustment will require not only a relatively high frequency of job creation and destruction at micro level but also sizeable job and worker flows.

## 2 Empirical strategy

Our analysis of employment adjustment at the *aggregate level* assumes a competitive, cost minimizing firm that takes output demand and input prices as given. In log-linear form, the (conditional) labor demand function can be expressed

$$l_t = \alpha_0 + \alpha_1 y_t + \alpha_2 w_t + \alpha_3 E_t + \beta T + u_t, \tag{1}$$

where $l$ denotes labor demand, $y$ is output demand, $w$ is the real wage, and $E$ is the input price of energy. This parsimonious specification also includes a deterministic trend term ($T$) to control for changes in employment not explained by output and input price growth. (For a similar treatment, see for example Flaig and Steiner 1989).[1] To be interpreted as a long-run – that is, cointegrating – relationship, the estimated residuals must be stationary, or I(0). Given the size of the sample (our data comprise quarterly observations over two decades), we will test this hypothesis within the framework of a single-stage error correction model (ECM), in which both the long-run relationship on the levels of the variables in Eq. (1) and the deviations from the long-run equilibrium are jointly estimated according to the following specification

$$\Delta l_t = \mu + \lambda \lfloor l_{t-1} - \alpha_1 y_{t-1} - \alpha_2 w_{t-1} - \alpha_3 E_{t-1} - \beta T \rfloor + \\ + \sum_{i=1}^{m} \delta_i \Delta l_{t-i} + \sum_{i=0}^{m} \gamma_{1i} \Delta y_{t-i} + \sum_{i=0}^{m} \gamma_{2i} \Delta w_{t-i} + \sum_{i=0}^{m} \gamma_{3i} \Delta E_{t-i} + e_t. \tag{2}$$

This procedure has thus the advantage of testing the long-run relationship between the levels of the relevant variables, while at the same time providing standard errors for the short-run and long-run effects (given by the parameters $\gamma$ and $\alpha$, respectively) and for the error correction coefficient, $\lambda$. Note that the statistical significance of the parameter $\lambda$ will serve as a cointegrating test, so that rejection of the null (that $\lambda$ is zero) will be interpreted as rejection of the long-run relationship in Eq. (1).[2] The parameter $\lambda$ also gives an aggregate measure of how firms react to past deviations from the assumed long-run equilibrium. However, given the over-parameterization of the model in Eq. (2), the dynamics of labor adjustment

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[1] Firm behavior may of course differ if countries operate distinct product market regimes. According to the OECD index of product market regulation, Germany appears less regulated than Portugal. The overall index for Germany is 1.4 as compared with 1.7 for Portugal. (The OECD average is 1.52, with a standard deviation of 0.47; see Nicoletti et al., 2000, Table A3.7.) However, cluster analysis indicates that the two countries belong to the same bloc of nations characterized by relatively restrictive/interventionist practices (Nicoletti et al., 2000, Figure 6).

[2] In testing the hypothesis that $\lambda$ in this model is statistically different from zero (i.e. that the variables are cointegrated), we will use the critical values reported by Banerjee, Dolado, and Mestre (1998).

are subsumed in all the lagged parameters, and not just in $\lambda$. To simplify the interpretation of the dynamic properties of the model, therefore, the estimated parameters will be used to simulate the impact of a one-time exogenous shock on employment adjustment.

The analysis of employment adjustment at the *firm level* follows a different strand of the literature. The framework is now one in which firms are assumed to respond to common (aggregate) demand shocks, as well as to firm-specific shocks (see, inter al., Arellano and Bond 1991; Bentolila and Saint-Paul 1992; Blundell and Bond, 1998). And although the time span of the firm panel series is much shorter than for the time series data, the richer information available at this level allows us to study the dynamics of employment adjustment in greater detail.

The dynamic specification of labor demand contains the input prices of labor and materials, the stock of capital, and one lagged employment term to control for sluggish labor adjustment. Changes in labor demand are also a function of specific and general demand shocks. The former are proxied by the (log) change in firm sales, and the latter by time dummies.

Formulated in logs, a general model for panel data can be written

$$l_{it} = \lambda l_{it-1} + \beta'(L)X_{it} + f_i + v_t + e_{it} \tag{3}$$

where $L$ is the lag operator, and $\beta$ is the vector of coefficients of exogenous variables. (As a practical matter, the input prices of labor and materials will be treated as endogenous variables, given that they are obtained by dividing total costs by total employment.) All unobservable variables specific to the individual firm are captured by the time-invariant firm-specific component $f_i$. Macroeconomic events (i.e. aggregate demand shocks) specific to a given year are represented by $v_t$. Finally, $e_{it}$ is a white noise residual.[3]

We note that the model in Eq. (3) can be obtained by assuming the conditional labor demand (with just one RHS variable, $x_{it}$ , for expositional convenience)

$$l_{it} = \beta_1 x_{it} + f_i + v_t + u_{it}, \tag{3.1}$$

and an AR(1) structure for the error term of the form

$$u_{it} = \lambda u_{it-1} + e_{it}. \tag{3.2}$$

After substitution and rearranging terms, we have

$$l_{it} = \lambda l_{it-1} + \beta_1 x_{it} - \beta_1 \lambda x_{it-1} + n_i^* + v_t^* + e_{it}, \tag{3.3}$$

which, under the 'common-factor' restrictions $-\beta_1\lambda = \beta_2$, $n_t^* = (1-\lambda)f_i$, and $v_t^* = v_t - \lambda v_{t-1}$, is equivalent to

$$l_{it} = \lambda l_{it-1} + \beta_1 x_{it} + \beta_2 x_{it-1} + f_i + v_t + e_{it}. \tag{3.4}$$

[3]The time-series analysis of aggregate data covers a period of two decades (1977–1997) of quarterly observations, while the panel of firms covers a maximum of nine periods of annual observations. Given the short length of the panel, the dynamic panel specification uses standard static conditional labor demand arguments together with a lagged employment regressor to account for sluggish labor adjustment. Although the results from specifications (2) and (3) are therefore not strictly comparable, they do provide an acceptable indication of the degree of employment inertia present in aggregate time-series and firm-level data.

If we set $\beta'(L)X_{it} = \beta_1 x_{it} + \beta_2 x_{it-1}$, equation (3) becomes (3.4). And it also follows that an AR(2) error structure would yield a similar autoregressive model, with additional $l_{it-2}$ and $x_{it-2}$ terms.[4]

Returning to Eq. (3), given the presence of lagged dependent variables on the right-hand side of the equation (together with the presence of endogenous input price variables), use of standard panel estimation techniques would produce biased and inconsistent estimates. We therefore employ the Generalized Method of Moments (GMM) estimator developed by Arellano and Bond (1991). This approach extends the first difference instrumental variables method suggested by Anderson and Hsiao (1981) to dynamic fixed-effects models, and yields asymptotic standard errors that are robust to general cross-section and time-series heteroskedasticity under the null hypothesis of no serial correlation in the errors. (To test the latter hypothesis, Arellano and Bond have developed a first- and second-order serial correlation test statistic based on the GMM residuals.) We will use in particular two alternative linear estimators: the GMM-DIF and the GMM-SYS (after Blundell and Bond 1998). The first method uses the standard Arellano-Bond first-differenced GMM estimator, while the second combines transformed (i.e. first-differenced) and levels equations. As shown by Blundell and Bond, the latter method is supposed to yield more precise parameter estimates and to reduce potentially important small sample bias arising from the typical panel short sample periods.

Finally, we complement our aggregate (time-series) and disaggregate (panel) analyses with evidence on job and worker flows. Job creation/destruction is defined as the sum of the positive/negative firm- or establishment-level employment changes between $t$ and $t-1$, expressed as a percentage of the average employment level. The observed employment changes will also be used to compute a firm's transition probabilities between the three regimes of zero, negative, and positive job creation. Our analysis of worker flows will focus on hires and separations, defined as the flows of workers into and out of establishments between $t$ and $t-1$. The (gross) job reallocation (or job turnover) is then given by summing job creation and job destruction (i.e. the sum of employment gains at expanding units and employment losses at shrinking units) and (gross) worker reallocation (or worker turnover) by summing hires and separations.

## 3 Data

In this section we provide a description of the panel information used in the present inquiry. A detailed discussion of the aggregate manufacturing data is contained in Addison and Teixeira (2001). Suffice it to say here that the time series for employment, output, the real wage, and the price of energy were taken from various national statistical sources and from the OECD Main

[4]Derivations (3.1) through (3.4) are intended to show that the dynamic employment equation (3) can be interpreted as an extension of a standard conditional labor demand function, with the dynamics arising from an AR(1) disturbance term (Blundell, and Bond, 1998). Given that our implementation of the autoregressive panel data model in equation (3) does not impose any 'common-factor' restrictions, our findings in section 4.2 may be thought as an approximation to a more general process of adjustment.

Economic Indicators database (quarterly series), and cover the interval 1977:1–1997:4.

For Portugal, the analysis of employment adjustment at firm level uses an initial sample of 1,970 firms taken from the balance sheet records (*Central de Balanços*) of the Bank of Portugal for the period 1990-97 (annual observations). This database contains information on sales, cost of materials, labor costs (in the form of the wage bill), fixed and intangible assets, and the level of total employment. We applied several filters to the original sample, resulting in a balanced panel of 1,552 manufacturing firms (excluding oil refining). All firms in the sample had to have at least five paid employees and both the volume of sales and the cost of materials were restricted to at least 1,000 *contos* (thousand escudos) a year. The input price of materials and labor were obtained by dividing the relevant total costs by the number of employees. All nominal variables are expressed in real terms (1995 prices) using the GDP deflator.[5]

Unfortunately the balance sheet records at the Bank of Portugal do not provide any information on worker flows. The analysis of these flows will have to be based on different sources. In particular, we will cite research by Varejão (2001) and Portugal (1998) who use data from unbalanced panels of continuing establishments for the periods 1991–95 and 1983–94, respectively.

Our German 'firm' data are taken from the Establishment Panel of the Institute for Employment Research (*Institut für Arbeitsmarkt- und Berufsforschung, IAB)* of the Federal Employment Agency (*Bundesagentur für Arbeit*). The basis of the IAB panel is the *employment statistics register* of the Federal Employment Agency. Each year, all employers are required to report levels of and changes in the number of their employees who are subject to the compulsory social security scheme. The Establishment Panel draws a stratified random sample of units from the register, the selection probabilities depending on the employment frequency of the respective stratum. The strata comprise some 16 industries and 10 establishment size intervals covering all sectors and employment levels (see Kölling 2000). We applied several filters to the IAB sample, excluding three sectors (agriculture, financial services, and insurance) and non-profit organizations. The exclusions resulted in an unbalanced longitudinal data set containing some 13,200 establishments over the sample period 1993–2001. From this sample we also extracted a balanced panel of 800 units.

Variables taken from the IAB Establishment Panel include information on employment (e.g. skill composition of the workforce and the number of workers under fixed-term contracts), sales, gross wages, and intermediate inputs. The panel also contains information on replacement investment (and expansion investment) permitting estimation of the capital stock through time but the large number of missing observations meant that we ultimately eschewed calculating this measure. All nominal variables were deflated by the

[5]We calculated the stock of capital in the manner of Bentolila and Saint-Paul (1992, p. 146). This procedure involves computing (a) the initial market value of the capital stock (tangible assets) for each firm at the beginning of the period (this calculation assigns a given average depreciation rate and proxies the price of capital by the deflator of gross fixed capital formation), and (b) the real capital stock for successive years, based on the initial market value computed in (a) plus investment minus annual depreciation.

GDP implicit price level (OECD data). Finally, information on hirings and firings (including those to and from fixed-term contracts) is available as well (on a semi-annual basis), and is also used in our analysis of worker flows.

## 4 The Process of employment adjustment

### *4.1 Findings from aggregate data*

Formal unit root tests of the relevant series/variables – output ($y$), employment ($l$), the real wage ($w$), and the price of energy ($E$) – over the sample period 1977:1–1997:4 are reported in the Appendix. The first column of the table shows the results from the ADF test on levels of the variables. The null hypothesis is that the series are integrated of order one at the zero frequency (i.e. in the long-run). In none of the cases does the $t_{ADF}$ statistic exceed (in absolute value) the critical value. To further check on the presence of a single unit root, we applied the Dickey-Pantula (1987) procedure, which first tests for two unit roots (the second column) and then for a single unit root (third column). Based on these tests, employment, output, the real wage, and the relative price of energy are rejected as I(2) but not as I(1).[6]

Accordingly, the proper way to proceed is through cointegration analysis, and to this end we will focus on a single-stage ECM model estimation procedure, as described in section 2, in which the short- and the long-run parameters are jointly estimated. Proceeding therefore with the (nonlinear) estimation of model (2), the fitted regressions for Portugal and Germany are given in Table 1.

Clearly, in the case of Portugal there is evidence of a well-determined process of adjustment to the long-run equilibrium: the error correction term is strongly significant (the critical value for the $t_{ECM}$-test at the 0.01 level is $-4.60$), large in absolute value (namely 0.475 or 47.5% per quarter), and of the expected negative sign. In contrast, Germany reveals a much slower rate of adjustment to disequilibrium: a rate of 10.8% per quarter.[7] In both cases, the regression statistics are within the expected range – and the coefficient estimates for the levels variables $y_{t-1}$, $w_{t-1}$ and $E_{t-1}$ – that is, the long-run elasticities – are all statistically significant.[8] However, the $\lambda$ parameter by no means exhausts the dynamics of labor demand because the dynamic adjustment is also reflected in the coefficients of the lagged differences

[6] Additional tests, including unit root tests with structural breaks (Zivot and Andrews, 1992) and computation of confidence intervals for the largest autoregressive root (Stock, 1991), are documented in Addison and Teixeira (2001).

[7] Recall that apart from its serving as a cointegrating test – rejection of the null (that $\lambda$ is zero) can be interpreted as rejection of no cointegration – the $\lambda$ parameter also informs us of how firms react to past deviations from the long-run equilibrium.

[8] As documented by Hamermesh (1993a, Table 7.5), cross-country estimates of long-run employment-output elasticities range from 0.03 to 0.98. In the case of Germany, three estimates are reported: 0.16, 0.44, and 0.98. Although no values are reported for Portugal, our estimates of 0.54 and 0.89 for Portugal and Germany, respectively, may be said to lie within an ‘expected’ range.

**Table 1.** Nonlinear single-stage ECM model

| Variables | Portugal | Germany |
|---|---|---|
| Constant | 1.716 | 1.316 |
| | (0.342) | (0.260) |
| $\lambda$ | −0.475 | −0.108 |
| | (0.075) | (0.019) |
| $y_{t-1}$ | 0.539 | 0.884 |
| | (0.072) | (0.148) |
| $w_{t-1}$ | −0.213 | −0.566 |
| | (0.072) | (0.338) |
| $E_{t-1}$ | −0.059 | −0.061 |
| | (0.012) | (0.031) |
| $T$ | −0.0008 | −0.0025 |
| | (0.0004) | (0.0009) |
| $\Delta l_{t-1}$ | 0.294 | 0.255 |
| | (0.102) | (0.117) |
| $\Delta l_{t-2}$ | 0.277 | 0.407 |
| | (0.105) | (0.105) |
| $\Delta y_t$ | | 0.100 |
| | | (0.016) |
| $\Delta y_{t-2}$ | −0.419 | |
| | (0.123) | |
| $\Delta w_{t-1}$ | | −0.075 |
| | | (0.039) |
| $\Delta E_{t-1}$ | | 0.008 |
| | | (0.008) |
| $R^2$ | 0.62 | 0.94 |
| SER | 0.0058 | 0.0021 |
| LM(4) | 2.39 | 0.49 |
| ARCH | 0.46 | 0.14 |
| NORM | 1.92 | 0.46 |
| WHITE | 1.11 | 1.0 |

*Notes*: The specification for the single-stage ECM model is given by equation (2) in the text. SER is the standard error of the regression (an estimator of the variance of the disturbance term); LM(4) is the fourth order autocorrelation test; ARCH is the test for autoregressive conditional heteroskedasticity; NORM is the Jarque-Bera test for the normality of the residuals; and WHITE is White's test for heteroskedasticity based on the squares of the regressors. Standard errors are given in parenthesis.

included in the model. Because of this, we next examine the dynamic adjustment properties of labor demand graphically. In this exercise, we simulate the impact of a once-for-all exogenous shock in the employment equation. We expect that greater flexibility should manifest itself in a faster rate of adjustment of the labor input to the long-run equilibrium.

The results of the simulation for our two countries are given in panels (a) and (b) of Fig. 1. Since the main goal is to estimate the speed with which employment converges towards its long-run equilibrium, we expand the first difference operator in Eq. (2) to obtain a dynamic labor demand equation in the levels of the variables. A once-for-all shock is then introduced in this equation to simulate the adjustment path of labor. Despite its reputation as an exemplar of terminal inertia in labor markets,

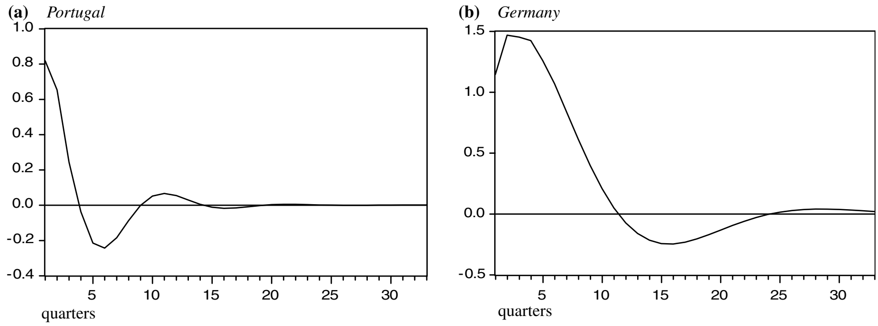


**Fig. 1.** Employment response to a one-time positive exogenous shock

note that Portugal shows a surprising speed of adjustment at this level of aggregation, with the changes in employment in response to the exogenous shock declining very quickly. On the other hand, the adjustment path observed for Germany seemingly confirms that changes in employment in response to an exogenous output shock follow a substantially slower adjustment path vis-à-vis Portugal.

### *4.2 Findings from firm-level data*

The fitted version of Eq. (3), in first differences, is presented in Table 2. Results for Portugal are given in row 1 of the table. In the case of Germany, we will present two sets of estimates: one based on a balanced sample of 800 establishments as for Portugal (in row 2), and the other drawn from a larger, unbalanced panel containing 1,632 establishments (row 3). For each country two alternative linear estimates – GMM-DIF and GMM-SYS – are provided. Year dummies are included in all specifications. As expected, the errors are negatively first order serially correlated, with no evidence of second order serial correlation. Further, the null of over-identifying restrictions (the Sargan/Hansen test) is not rejected, and the joint insignificance of the coefficients included in the regression is clearly rejected by the Wald test. Finally, the long-run employment elasticities with respect to input prices (labor cost and the price of materials/intermediate input) and output changes are in general smaller than their time-series counterparts, but are in the usual panel estimates range.

Using the GMM-DIF method, the coefficient estimate for the lagged employment variable in Portugal is 0.77. Employment adjustment is clearly higher than in Germany where the corresponding coefficient is 0.87 (or 0.83 for the unbalanced sample). Taken together, the evidence provides a consistent ordering of employment adjustment, with Germany evincing the greater degree of employment inertia despite its more favorable reputation. The GMM-SYS coefficient estimates maintain this 'ranking' but, as expected, they are substantially higher. This is especially so for Germany, where the estimated degree of employment inertia is an extremely high 0.97. That said, we caution that there is overlap of the 95% confidence intervals in

**Table 2.** First difference estimates of labor demand in Portugal and Germany

| Country/ Sample | Data set | Variable | Model GMM-DIF | GMM-SYS |
|---|---|---|---|---|
| 1. Portugal (1990–1997) | Balanced panel of manufacturing firms. | $l_{it-1}$ | 0.770 (0.050) | 0.858 (0.044) |
| | Database: Central de Balanços, Bank of Portugal. | $w_{it}$ | −0.486 (0.117) | −0.576 (0.116) |
| | Number of units: 1,552 firms. | $w_{it-1}$ | 0.370 (0.043) | 0.452 (0.051) |
| | | $p_{it}$ | −0.273 (0.050) | −0.255 (0.046) |
| | | $p_{it-1}$ | 0.221 (0.037) | 0.250 (0.037) |
| | | $k_{it}$ | 0.163 (0.049) | 0.107 (0.035) |
| | | $shock_{it}$ | 0.467 (0.060) | 0.441 (0.060) |
| | | $m_1$ | −10.71 | −11.66 |
| | | $m_2$ | 2.10 | 1.97 |
| | | Wald | 1377 | 22120 |
| | | Sargan (df) | 70.48 (36) | 86.47 (56) |
| 2. Germany (1993–2001) | Balanced panel of economy-wide establishments | $l_{it-1}$ | 0.868 (0.160) | 0.971 (0.017) |
| | Database: IAB Establishment Panel. | $w_{it}$ | −0.094 (0.222) | 0.096 (0.075) |
| | Number of units: 800 establishments. | $w_{it-1}$ | 0.022 (0.038) | 0.026 (0.033) |
| | | $p_{it}$ | −0.047 (0.096) | −0.157 (0.068) |
| | | $p_{it-1}$ | 0.045 (0.082) | 0.147 (0.051) |
| | | $shock_{it}$ | 0.312 (0.149) | 0.302 (0.115) |
| | | $m_1$ | −3.36 | −5.24 |
| | | $m_2$ | 0.44 | 0.87 |
| | | Wald | 55.85 | 5884 |
| | | Sargan (df) | 5.89 (12) | 26.43 (32) |
| 3. Germany (1993–2001) | Unbalanced panel of economy-wide establishments | $l_{it-1}$ | 0.834 (0.212) | 0.967 (0.018) |
| | Database: IAB Establishment Panel. | $w_{it}$ | 0.322 (0.234) | 0.069 (0.084) |
| | Number of units: 1,632 establishments. | $w_{it-1}$ | −0.047 (0.043) | −0.0005 (0.034) |
| | | $p_{it}$ | −0.059 (0.115) | −0.177 (0.071) |
| | | $p_{it-1}$ | 0.106 (0.088) | −0.157 (0.049) |
| | | $shock_{it}$ | 0.161 (0.233) | 0.422 (0.154) |
| | | $m_1$ | −3.37 | −8.27 |
| | | $m_2$ | 1.32 | 1.95 |
| | | Wald | 42.68 | 12990 |
| | | Sargan (df) | 4.14 (12) | 37.96 (32) |

*Notes*: The estimates were obtained using the general model specified in equation (3) in the text (in first-differences). The estimated models include the stock of capital, $k$, firm-specific demand shocks, *shock*, the price of labor (labor cost), $w$, and materials/intermediate inputs, $p$. Due to data problems, the estimations for Germany do not include any measure of the capital stock. Asymptotic standard errors robust to general cross-section and time-series heteroskedasticity are given in parentheses. $m_1$ and $m_2$ are tests for first- and second-order serial correlation in the first-differenced residuals, *Wald* is a test of joint significance of the independent variables, and *Sargan* is a test of overidentifying restrictions. In the estimation, we have used the DPD software for OX, version 2.1, available at http://www.nuff.ox.ac.uk/Users/Doornik.
*Sources*: Central de Balanços of the Bank of Portugal; IAB Establishment Panel.

the GMM-DIF case. The standard errors are smaller in the GMM-SYS case, yielding slight overlap at 95% and none at 90%.[9]

Yet the panel estimates do seem to generate longer adjustment paths than their time-series counterparts. According to our time-series formulation, given an exogenous shock in, say, output demand, (aggregate) employment will react faster, the larger is the ECM term (viz. the $\lambda$ parameter) in Eq. (2). In turn, for the panel the reaction of firms' labor demand to an exogenous shock will be lower, the greater is the estimated coefficient of the lagged employment term in Eq. (3). One quick way of comparing employment inertia across these two tiers of data, therefore, is simply to connect the two. (Abstracting from the lagged differenced employment terms and all other right-hand side variables, it follows from Eq. (2) that $l_t = \mu + (1 + \lambda) l_{t-1} - \lambda \alpha_1 y_{t-1} + e_t, \lambda < 0$, while in Eq. (3) we have $l_{it} = \lambda l_{it-1} + \beta'(L) X_{it} + f_i + v_t + e_{it}, \lambda < 1$.) Clearly, employment inertia captured in the time-series data is much lower than in the panel of firms, especially for the GMM-SYS estimates: for Portugal, the lagged employment coefficient in the time-series is approximately 0.53 (= 1 − 0.47), while in the panel it is equal to 0.86 in the case of Portugal; for Germany, the corresponding values are 0.89 (= 1 − 0.11) and 0.97, respectively.[10]

With respect to Portugal, it will be recalled that the panel estimates of the speed of adjustment are drawn from a balanced sample of surviving (manufacturing) firms, in which there is an over-representation of large units and no consideration of the entry and exit of firms. Accordingly, the panel estimates could reflect the failure to control for the mobility of firms.[11] That being said, the two sets of estimates provided for Germany perhaps indicate that the bias introduced by using panels of continuing/surviving firms is of

[9]For Portugal, the 95/90 percent confidence intervals (GMM-DIF) are 0.672 to 0.868/0.688 to 0.852. The corresponding intervals for Germany (balanced panel) are 0.554 to 1.182/0.605 to 1.131 (0.418 to 1.250/0.485 to 1.183, unbalanced panel). For GMM-SYS, the 95/90 percent confidence intervals are 0.772 to 0.944/0.786 to 0.930 for Portugal; in Germany (balanced panel) they are 0.938 to 1.004/0.943 to 0.999 (0.932 to 1.002/0.937 to 0.997, unbalanced panel).

[10]For the purposes of illustration, if we further assume a partial adjustment mechanism of the Koyck type, it follows that the time-series relationship $l_t = \mu + (1 + \lambda) l_{t-1} - \lambda \alpha_1 y_{t-1} + e_t$ for Portugal yields a mean adjustment lag of $(1 + \lambda)/(-\lambda) = 0.525/0.475$, or 1.1 quarters, whereas the mean adjustment lag using the panel (GMM-DIF) will be $\lambda/(1 - \lambda) = 0.77/0.23$, or 3.3 years. The corresponding mean lags for Germany (balanced panel) are 8.3 quarters and 6.6 years, respectively. Although the Koyck transformation is a crude simplification – the unrestricted lag structures contained in equations (2) and (3) are rooted in a much more general dynamic structure – it offers a rough indication of the degree of employment inertia in the two types of data used here.

[11]Nevertheless, the relatively higher representation of small firms in the Portuguese economy will tend to imply larger (smaller) time-series (panel) estimates of employment adjustment vis-à-vis Germany. A different issue has to do with the role of specific categories of labor, namely, atypical (fixed-term) versus regular (or open-ended) employment. In the context of autoregressive panel data models, the latter question has been addressed by examining labor demand stability, testing whether the elasticity of employment to changes in output shocks is sensitive to the cycle (e.g. Bentolila and Saint Paul, 1992). We do not report any results using this modeling strategy, mainly because our Portuguese data set does not distinguish the labor input by type of contract (or wage), making the test as to the role of atypical work too indirect. The share of fixed-term contract workers in Germany is also too small to make valid inferences.

secondary importance. As can be seen from row 3 of the table, use of the unbalanced panel (i.e. allowing for plant accessions and exits) generates only slightly lower employment inertia than does use of the balanced sample. Furthermore, this result obtains irrespective of the estimation method. Nevertheless, two caveats are in order. First, panel estimates in first differences impose strong data requirements, since each establishment must remain in the panel for at least three or four consecutive periods. Second, the process of adjustment is expected to be less dependent on plant openings and closings in Germany than in Portugal. Thus, for example, the contribution of such openings and closings to total job turnover for Germany is roughly one-half that for Portugal (OECD 1996; Portugal 1999).

We note finally that the time-series estimates in subsection 4.1 are subject to aggregation bias of uncertain magnitude. In most cases, aggregation over single units tends to produce a smoother pattern of adjustment (i.e. shorter lag lengths), such that the estimated speed of adjustment from Eq. (2) might bear little correspondence to underlying microeconomic behavior. In Portugal, the contrast between the very rapid speed of adjustment observed in the time-series data and the much slower pattern of adjustment found for the panel data does rather fuel this supposition. That said, our panel estimation is not without blemish. After all, we are using annual data that might be expected to introduce some *upward* bias and lead to greater employment persistence. Moreover, our procedure assumes common regressors across individual units, which can also bias upwardly the results (see Robertson and Symons 1992). We will now see whether examination of job and worker reallocation assists in understanding the advantages and limitations of each type of evidence.

### *4.3 Job and worker flows*

We seek in this section to identify the magnitude of job and worker reallocation that underlies the above estimates of employment persistence. As mentioned earlier, we anticipate that a smooth (or healthier) process of adjustment will require sizeable job and worker flows.

We will first comment on job reallocation based on disaggregate job flows (i.e. annual rates of job creation and job destruction at firm/establishment-level) obtained from the Portuguese and German panels. These annual averages are given in Table 3 (rows, 1, 4 and 5). For the Portuguese balanced panel of surviving firms, the job creation and destruction rates are 3.3 and 5.2%, respectively, yielding a job turnover rate of 8.5% (row 1). Job creation and destruction averages for the corresponding German sample (row 4) are clearly smaller at 1.8 and 4.6%, respectively. In other words, the German job turnover rate is more than two percentage points lower than in Portugal. The impression that job turnover in Germany is low is only reinforced by the fact that the German sample includes all sectors of the economy and not just manufacturing. As is well known, (excluded) sectors such as the retail trade and construction have among the highest turnover rates.

Fig. 2 and 3 provide additional information on job creation and job destruction in the Portuguese and German balanced panels. Fig. 2 charts the distribution of these job flows (respectively, job creation and destruction). In

**Table 3.** Average job and worker flows and transition rates in Portugal and Germany

| Country (sample)/Study | Data set | | Job flows | | | | | Worker flows | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | Job creation and destruction | | | Job turnover | | H | S | Worker turnover | |
| | | | JC | JD | $\Delta E$ | Quarterly | Annual | | | Quarterly | Semi-annual |
| 1. Portugal (1990–1997). Authors' own estimates. | Balanced panel of manufacturing firms. Database: Balance sheet records (*Central de Balanços*, Bank of Portugal). Number of units: 1,552 firms. | All | 3.3 | 5.2 | −1.9 | | 8.5 | | | | |
| 2. Portugal (1991:1–1995:4). Varejão (2001). | Unbalanced panel of continuing establishments. | PC | 1.9 | 2.6 | −0.7 | 4.5 | | 1.9 | 3.2 | 5.1 | |
| | Database: Economy-wide data (*Inquérito ao Emprego Estruturado*), | FTC | 9.8 | 12.0 | −2.2 | 21.8 | | 16.4 | 14.2 | 30.6 | |
| | Ministry of Labor); the primary sector and public administration were excluded. | All | 2.3 | 3.1 | −0.8 | 5.4 | | 4.0 | 4.8 | 8.8 | |
| | Number of units: 10,673 establishments. | All | 5.7 | 8.0 | −2.3 | | 13.7 | | | | |
| 3. Portugal (1983–1994). Portugal (1999). | Database: *Quadros de Pessoal* and *Inquérito ao Emprego Estruturado*, Ministry of Labor: | | | | | | | | | | |
| | Economy-wide | All | 14.9 | 13.7 | 1.2 | | 28.6 | | | | |
| | Economy-wide (excluding openings and closings) | All | 7.1 | 7.3 | −0.2 | | 14.4 | | | | |
| | Economy-wide | All | 4.0 | 3.9 | 0.1 | 7.9 | | | | | |
| | Economy-wide (excluding openings and closings) | All | 2.2 | 2.8 | −0.6 | 5.0 | | | | | |

**Table 3.** (Contd.)

| Country (sample)/Study | Data set | | Job flows | | | | | Worker flows | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | Job creation and destruction | | | Job turnover | | H | S | Worker turnover | |
| | | | JC | JD | $\Delta E$ | Quarterly | Annual | | | Quarterly | Semi-annual |
| 4. Germany (1993–2001). Authors' own estimates. | Balanced panel of continuing establishments. | PC | 1.4 | 4.0 | −2.6 | | 5.4 | | | | |
| | Database: IAB Establishment Panel. | FTC | 29.9 | 23.9 | 6.0 | | 53.8 | | | | |
| | The primary sector, insurance and banking, and non-profit organizations were excluded. | ALL | 1.8 | 4.6 | −2.8 | | 6.4 | 2.9 | 4.2 | | 7.1 |
| 5. Germany (1993–2001). Authors' own estimates. | Unbalanced establishment data. | PC | 2.3 | 3.5 | −1.2 | | 5.8 | | | | |
| | Database: IAB Establishment Panel. | FTC | 28.7 | 20.4 | 8.3 | | 49.2 | | | | |
| | The primary sector, insurance and banking, and non-profit organizations were excluded. | ALL | 2.8 | 4.9 | −2.1 | | 7.7 | 3.5 | 5.6 | | 9.1 |

*Notes*: JC, JD, H, S, and $\Delta E$ denote job creation, job destruction, hires, separations, and net employment change, respectively (measured either at quarterly, semi-annual or annual frequencies). We are following here the definitions given by Davis and Haltiwanger (1992). Thus, denoting employment in firm/establishment $i$ in year (quarter) $t$ by $l_{it}$, there is job creation (job destruction) if $l_{it} > l_{it-1}$ ($l_{it} < l_{it-1}$). In other words, job creation (job destruction) is equal to positive (negative) firm/establishment-level employment changes. Ignoring openings (closings) and summing up job creation (job destruction) across expanding (contracting) units gives the economy's job creation (destruction) rate, respectively JC and JD. (All measures are expressed as a percentage of average employment. Aggregate net employment change, $\Delta E$, is given by the difference between JC and JD and job turnover is the sum of JC and JD. Hires in unit $i$ in year (quarter) $t$ is given by the number of newly admitted workers during period $t$, while separations is given by the number of workers who left the firm/establishment during that period (again measured in terms of average employment). Worker turnover is given by hires plus separations. ALL, FTC, and PC denote total employment, employment under fixed-term contracts, and employment under open-ended contracts, respectively.

*Sources*: Central de Balanços of the Bank of Portugal; IAB Establishment Panel; Portugal (1998); Varejão (2001).

24% of the cases there is neither job creation nor job destruction in Germany; for Portugal, the proportion is only 19%. *Vulgo*: Portuguese firms tend to adjust labor input more often. In a further 45% of cases in both countries, job creation and job destruction rates do not exceed 10% (in absolute value). This finding indicates that the dispersion of job flows at micro level in either country is too large to admit of a thoroughly lumpy process of employment adjustment. Were that the position, we should observe a greater concentration of job flows (in either direction) in firms with sharp employment changes.

Job creation and destruction can be further examined by constructing a hierarchy or ranking of observed annual job flow rates. That is to say, we can order or rank from highest to lowest (most negative in the case of job destruction) a firm's job flow rate for each year of the sample period and then take the mean rate across units for each 'rank.' There are seven such rankings in all given the seven-year sample periods. The results of this exercise are given in Fig. 3. Reading from right to left, the height of the first (second, third, etc.) column(s) denotes the mean of the largest (second largest, third largest, etc.) job flow rate estimated over all units in the sample. The goal is to compare these grand averages or, equivalently, differences in the height of the seven columns. The smoother the process of employment adjustment, the smaller these differences should be. For example, the second column (denoted by 'rank 2') has roughly one-half the height of the first column (rank 1) and a greater difference would signal more lumpy employment adjustment. In turn, while it is true that in both countries large job creation/destruction flows are generally followed by smaller flows in subsequent periods, the observed pattern is again suggestive of a smoother process of employment adjustment in Portugal by virtue of there being smaller differences across columns than is the case for Germany.

We can also use firm/establishment-level job flows to measure the probability of a unit's shifting between job creation and job destruction 'regimes,' where regime 1 will signify an absence of either job creation or job destruction and regimes 2 and 3 will respectively characterize job creation

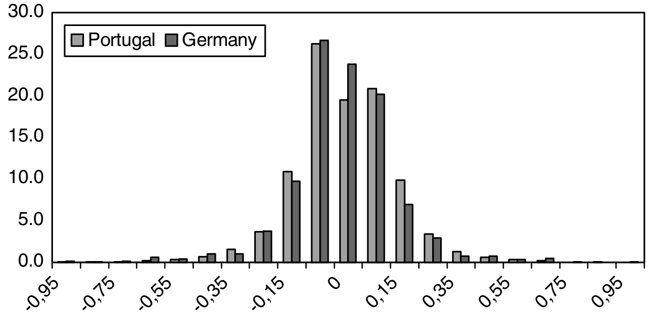


**Fig. 2.** Distribution of job flows (job creation and destruction) in Portugal and Germany (percent)

*Notes*: The 10-point intervals on the x-axis are represented by the corresponding midpoints. The sample periods are 1990–1997 (Portugal) and 1993–2001 (Germany).

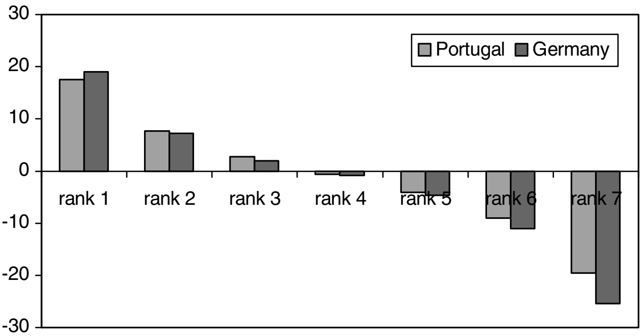


**Fig. 3.** Ranking of job flows in Portugal and Germany (percent)
*Notes:* See the text for the derivation of the seven ranks. The height of columns 1 through 7 denotes the average creation/destruction rate of each rank. The sample periods are 1990-1997 (Portugal) and 1993–2000 (Germany).

and job destruction. Table 4 presents the respective transition rates. In the case of Portugal, the probability of remaining in regime 1 in any two consecutive periods (0.34) is not large enough to preclude a fairly smooth process of adjustment. This tendency is further confirmed by the values of the second and third elements of the principal diagonal (0.40 and 0.50, respectively), and by the low probability of transitioning out of regimes 2 or 3 into 1 (0.16 in each case). For Germany as well the second and third elements of the principal diagonal are also too large to diagnose terminal arteriosclerosis, although some indication of that country's greater employment inertia is signaled by the much higher probability of remaining in regime 1 (namely, 0.48).

Both of the above pieces of evidence accord with the relative positions of the two countries reported in the panel estimates. Furthermore, the quarterly job turnover rate of 5.0–5.4 percent (Table 3, rows 2 and 3) suggests that there are nontrivial micro-level employment changes underlying the relatively lower degree of employment inertia detected in the Portuguese aggregate (quarterly) time-series. Note also that the higher turnover value of 7.9% – obtained when we take account of openings and closings in row 3 – suggests that we should indeed expect a lower speed of adjustment in panel estimates of continuing establishments/firms.[12]

Finally, Table 3 also contains some summary measures of worker flows. Since similar job flows can hide substantial differences in worker flows, it is worth briefly examining this possibility. We have shown that a more rapid adjustment of employment to output changes in Portugal (in both time-series and panel estimates) finds expression in a higher rate of job turnover. Is this also reflected in higher worker mobility? Rows 2 and 4 of

[12]Disaggregation by type of employment contract shows that fixed-term contract (FTC) workers in Portugal do play a role in the relatively high speed of employment adjustment observed in quarterly time-series. As can be seen in Table 3, the Portuguese quarterly job turnover rate of FTC workers of 21.8 percent (row 2) is approximately one-half the annual figure for Germany (rows 4 and 5).

**Table 4.** Annual transition probabilities between job creation and job destruction regimes in Portugal and Germany

| | Regime 1 Portugal | Germany | Regime 2 Portugal | Germany | Regime 3 Portugal | Germany |
|---|---|---|---|---|---|---|
| Regime 1 | 0.34 | 0.48 | 0.30 | 0.25 | 0.36 | 0.27 |
| Regime 2 | 0.16 | 0.17 | 0.40 | 0.43 | 0.44 | 0.40 |
| Regime 3 | 0.16 | 0.15 | 0.34 | 0.31 | 0.50 | 0.54 |

*Notes*: The transition rate in cell $P_{ij}$ is the probability of a firm/establishment shifting between regime i to regime j over the course of a year, i, j = 1, 2, 3. In Regime 1 there is no job creation or destruction; Regime 2 (Regime 3) characterizes job creation (job destruction). The sample periods are 1990–97 (Portugal) and 1993–2001 (Germany), annual observations.

Table 3 indicate that this is indeed the case. Thus, in the case of Portugal we observe a quarterly worker turnover rate of 8.8%, while in Germany the semi-annual worker turnover rate is the range 7 to 9% (rows 4 and 5), or roughly half the corresponding Portuguese value of 17.6% (i.e. $2 \times 8.8\%$).

## 5 Summary

This paper has provided a layered perspective of the process of labor adjustment in two countries with unenviable labor market reputations. It has combined three different types of information: aggregate time-series data, firm/establishment panel data, and job and worker flow data. Although evidence derived from any one layer is insufficient to sustain strong conclusions about the process of labor market adjustment, we found consistencies in the evidence from all three layers, which permits a more thorough assessment of country 'reputations' and of differences between countries.

We began our discussion with an analysis of aggregate time-series data for manufacturing and applied a single-stage error correction model to study the behavior of employment. The main advantage of this approach was that it enabled us to chart the adjustment process over a reasonably long interval. For our two countries we detected material differences in the speed of adjustment of employment to deviations from the long-run labor demand equilibrium. This was demonstrated not only by the magnitude and statistical significance of the error correction terms but also through a simulation exercise that accounted for all the parameters of the estimated dynamic model. The exercise proved favorable to Portugal.

Aggregate analysis, however useful, does not exhaust the estimation options. In response, we first analyzed labor adjustment at the firm/establishment level, using information from two original micro data sets. Although the frequency of our panel data was such as to preclude investigation of the within-year responses of firms and establishments to output shocks, the size and reach of the samples and the richness of the explanatory variables enabled us to offer a reasonably complete treatment of labor demand at the micro level. Our estimates of employment inertia were consistent with the aggregate evidence; that is, we reported evidence of a lower coefficient estimate for lagged employment for Portugal, even if the 95% confidence intervals did overlap.

We noted that our firm/establishment data are not without blemish. In particular, the annual frequency of the panels may artificially inflate the lags in labor adjustment, and our estimates of employment inertia should also be higher when using balanced panels. Thus, in the next and final stage of this empirical inquiry, we exploited some descriptive data on job and worker flows for the two countries. The goal was not so much to offer a different perspective on the process of labor adjustment as it was to hopefully strengthen the credibility of the conclusions drawn from the econometric estimates. We reported evidence in favor of a higher frequency of employment changes and more sizeable job and worker flows in Portugal.

Putting information from all three tiers of data in to perspective is a nontrivial task. But the broad lesson of a reasonably benchmarked exercise is that one of the countries, Portugal, seems on closer inspection to have been unfairly indicted in policy councils as the exemplar of labor market rigidities. That said, the reputation of the other country, Germany, is not improved by this layered examination of its process of employment adjustment.

We conclude by noting some limitations of our approach. In the first place, our analysis has focused on the adjustment of employment to output and, as is conventional, we have restricted attention to the demand side of the labor market and abstracted from supply-side considerations. The results reported here are therefore conditional on the partial equilibrium nature of our modeling strategy. Second, our labor demand specification relates solely to the demand for employment (workers), so that we have ignored the role of hours in the adjustment of (total) labor input. In the case of Germany, there is plenty of evidence to suggest that firms have been able to adjust fairly rapidly along the hours margin in response to output shocks (see, for example Abraham and Houseman 1994; Hunt 2000). That said, while a high speed of adjustment in hours will reflect high costs of employment adjustment, it will likely also reflect suboptimal adjustment of labor input to exogenous shocks. In other words, success in circumventing the restrictions imposed by high employment protection through hours adjustment – that is, more flexible use of hours and short time work – may not be profitable. In addition, the fact that Germany exhibits a low index of hours worked might well suggest that adjustment along the hours dimension is close to exhaustion.[13] A related final issue is the level of aggregation of the labor input in this treatment. We have not disaggregated by gender and skill, nor have we analyzed the influence of part-time work versus full-time employment on the process of labor adjustment. In each case, investigation of these issues is the stuff of a future research agenda.

[13]Indeed, Germany evinces one of the lowest number of hours worked per person of working age, ranking 14 out of 20 OECD countries; Portugal, by contrast, ranks third highest (see Nickell and Nunziatta, 2000, Table 5).

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## Appendix

Appendix: Augmented Dickey-Fuller (ADF) and Dickey-Pantula Univariate Tests

| Series | Augmented Dickey-Fuller | | | Dickey-Pantula | | | Dickey-Pantula | | |
|---|---|---|---|---|---|---|---|---|---|
| | Ho: $y_t \sim I(1)$<br>H1: $y_t \sim I(0)$ | | | Ho: $y_t \sim I(2)$<br>H1: $y_t \sim I(1)$ | | | Ho: $y_t \sim I(1)$<br>H1: $y_t \sim I(0)$ | | |
| | t | lags | F(4, T) | t | lags | F(4, T) | t | lags | F(4, T) |
| Portugal*l* | | | | | | | | | |
| *l* | −1.38 | 1,2,3,5 | 1.24 | −6.33 | 1,2,4 | 1.07 | −1.68 | 1,3,4 | 1.55 |
| *y* | −2.75 | 2,3 | 1.20 | −4.49 | 1 | 0.69 | −2.29 | 1 | 1.53 |
| *w* | −1.29 | 1,4 | 1.96 | −4.21 | 1,3 | 1.57 | −2.43 | 4 | 0.73 |
| *E* | −1.56 | 1 | 0.89 | −4.54 | 1 | 1.10 | −1.50 | 1 | 1.22 |
| Germany | | | | | | | | | |
| *l* | −2.00 | 1,2,4 | 0.35 | −3.45 | 4,5,6 | 1.66 | −1.87 | 1,2,3 | 0.14 |
| *y* | −2.55 | 2,3 | 1.10 | −7.96 | 3 | 0.28 | −1.78 | 3 | 0.82 |
| *w* | −3.26 | 1-7 | 0.93 | −9.03 | 7 | 1.24 | −2.54 | 7 | 1.28 |
| *E* | −3.33 | 1-9 | 0.37 | −4.03 | 1,2,3 | 0.39 | −2.76 | 2,3 | 1.26 |

*Notes*: The ADF equation is $\Delta z_t = \alpha + \beta_0 z_{t-1} + \sum_{i=1}^{k} \beta_i \Delta z_{t-i} + \delta T + u_t$, and the null hypothesis is that the series are not stationary: $\beta_0 = 0$. The lag structure is such that the errors are white noise. The F(4, T) statistic tests for the presence of fourth order serial correlation in the residuals of the ADF equation (where the null is absence of autocorrelation). MacKinnon critical values for the ADF test are −4.04 and −3.45 at 1% and 5%, respectively. The auxiliary equations for the Dickey-Pantula test of Ho: I(2) against H1:I(1), and Ho: I(1) against H1: I(0) are, respectively, $\Delta(\Delta z_t) = \alpha_0 + \alpha_1 \Delta z_{t-1} + \sum_{i=1}^{k} \beta_i \Delta(\Delta z_{t-i}) + \delta T + u_t$, and $\Delta(\Delta z_t) = \alpha_0 + \alpha_2 \Delta z_{t-1} + \alpha_3 \Delta z_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta(\Delta z_{t-i}) + \delta T + u_t$.
The t-values of the coefficient estimates ($\alpha_1 = 0$ and $\alpha_2 = 0$, respectively) follow a non-standard DF distribution. *l* denotes employment, *y* output, *w* the real wage, and *E* the price of energy (in logs). The sample period is 1977:1–1997:4.