Macroeconomic Dynamics, **2**, 1998, 472–491. Printed in the United States of America.

WAGE ADJUSTMENT AND EMPLOYMENT PERSISTENCY

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Persistent high unemployment has fostered the employment persistency hypothesis according to which employment changes are driven by unanticipated shocks whereas anticipated shocks that potentially could change employment are absorbed by wage changes. Empirical tests of the persistency hypothesis fail to distinguish between the properties of shocks and endogenous propagation mechanism causing persistency. This paper develops a new test strategy by explicitly distinguishing between these two factors. The methodology is applied to the manufacturing sector in Denmark, and some support in favor of an endogenous propagation mechanism causing employment persistency is found.

Keywords: Insider-Outsider Model, Wage Employment Error Correction Model

1. INTRODUCTION

The unemployment rate has displayed persistency in a number of countries over the past couple of decades, and this remain a serious concern for economic policy. Persistency in employment and thus unemployment[1] can arise either because shocks are persistent or because adjustment mechanisms cause even temporary shocks to have permanent effects. It is important to distinguish between these two sources of persistency because the latter implies that unemployment cannot be reduced by policies directed at increasing labor demand. Existing empirical analyses fail to distinguish between these two sources of persistency and thus may provide little information of relevance for designing an effective economic policy toward the unemployment problem. This paper develops a test strategy to distinguish between the two sources of persistency by explicitly considering the adjustment of wages and employment to shocks.

A simple prototype insider-outsider model is used as a benchmark for developing the test strategy.[2] The model is a convenient vehicle for bringing out the type of wage adjustment that is needed for employment persistency to arise. The purpose here is not to identify empirically the possible sources of insider powers, but rather to clarify possible mechanisms leading to persistency in employment.

In addressing the adjustment of employment and wages to shocks, it is essential to take explicit account of the institutional structure of the labor market

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 1365-1005/98 $9.50

implying that wages are preset. Accordingly, wages and employment may respond differently to anticipated and unanticipated changes in labor demand. Persistency in employment only requires that anticipated changes in employment be reflected in wages so as to leave employment unaffected. Distinguishing between anticipated and unanticipated shocks yields as a side product insight on the importance of a contractual structure with preset wages, something that often is attributed central importance in macroeconomic analysis.

The test strategy that is developed is demanding in terms of the number of estimations needed. Hence, it is most transparent if carried out on a single-country basis. We have chosen Danish data because Denmark often is mentioned as an example of employment persistency and because this is the country with which we are most familiar.[3]

The paper is organized such that Section 2 outlines a simple prototype insider-outsider model that is a benchmark for formulating different test strategies for employment persistency. Section 3 tests the random-walk implication of the insider-outsider model, and Section 4 sets up an employment model that is used both to test for the so-called efficient-market property of employment and to identify the state variables for the wage-employment model in Section 5 that explicitly distinguishes between anticipated and unanticipated shocks. Section 6 offers some concluding remarks.

2. PROTOTYPE MODEL OF EMPLOYMENT PERSISTENCY

To develop a strategy for testing for employment persistency, a prototype insider-outsider model is used.[4]

Let the structural labor demand function (l_t) be given as (all variables in logs)

$$l_t = \alpha_0 + \alpha_1 w_t + \alpha_2 \mathbf{z}_t + \alpha_3 l_{t-1}, \qquad \alpha_1 < 0, \qquad 0 < \alpha_3 < 1. \tag{1}$$

Here, w_t is the real-product wage and $\mathbf{z}_t$ is a vector of state variables. Lagged employment enters to capture sluggishness on the part of firms in adjusting employment.

Consider a wage-setting scheme according to which the real product wage (effectively assuming full indexation of nominal wages) prior to each period is set so as to ensure employment of all currently employed, i.e., effective union membership or the insider group made up of those being employed in the last period [Blanchard and Summers (1986), Alogoskoufis and Manning (1988), and Blanchard and Fischer (1989)]. This implies that wage setting fulfils

$$\alpha_0 + \alpha_1 w_t + \alpha_2 E(\mathbf{z}_t \mid I_{t-1}) + \alpha_3 l_{t-1} = l_{t-1}, \tag{2}$$

and the wage equation reads

$$w_t = \alpha_1^{-1} [(l - \alpha_3) l_{t-1} - \alpha_0 - \alpha_2 \, E(\mathbf{z}_t \mid I_{t-1})], \tag{3}$$

https://doi.org/10.1017/S1365100598009043 Published online by Cambridge University Press

where $E(\mathbf{z}_t \mid I_{t-1})$ is the conditional expectation of the state variables conditional on the information set I_{t-1}, available to wage setters at the end of period $t-1$.

Inserting (3) into (1) yields the following reduced form for employment:

$$l_t = l_{t-1} + \alpha_2[\mathbf{z}_t - E(\mathbf{z}_t \mid I_{t-1})]. \tag{4}$$

It is an implication of the model that anticipated changes in labor demand are fully reflected in wages leaving employment unchanged, whereas unanticipated changes in labor demand affect employment but not wages.

Note that the model has a New Classical flavor in the sense that only unanticipated shocks can affect the level of employment and thus activity. Actually, the implications of the insider model are stronger than those of New Classical models because the ineffectiveness result applies not only to nominal shocks but also to real shocks. It is around this property of the model that the important policy conclusions arise.

This simple model is sufficient to outline different strategies for testing whether persistency arises as a result of the process for wage and employment determination. These are listed here in order of increasing generality, in the sense of being robust to reasonable modifications of the basic model. The innovative aspects of this paper lie primarily in the analysis based on a distinction between anticipated and unanticipated shocks.

2.1. Random-Walk Property of Employment

One test of the insider model arises directly from the reduced-form employment equation (4) according to which employment follows a random walk provided that the expectation errors are white noise.[5] The common way of testing this property proceeds by testing whether employment is integrated of order one, i.e., employment needs to be differenced once to become stationary. Although integration of order one is a necessary condition for the insider-outsider model to be validated, a test for integration is a weak test of the model.

First, persistency in employment can arise for many reasons. Most models of the labor market, including the competitive model, also predict that employment is highly persistent if the state variables are highly persistent. Hence, this test procedure is a useful first step in clarifying whether employment is highly persistent, but it does not shed light on whether this persistency arises because of persistency in market conditions or because of endogenous propagation mechanisms such as insider power.

Second, finding that employment is integrated of order one is a necessary but not sufficient condition for employment to follow a random walk, because a random walk is an integrated proces with a white-noise error term. The white-noise error property can be tested directly, but (4) also implies that changes in employment cannot be predicted on the information set I_{t-1}. That is, employment changes should have an efficient-market property contingent on information available at

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the time of the wage setting. A stronger test of the persistency model is, therefore, to test whether signals in the information set I_{t-1} help predict employment changes. If so, the efficient-market property of employment changes is invalidated.

2.2. Negative Effect of Lagged Employment on Wages

According to the wage equation (3), lagged employment affects wages negatively. This effect arises as the net effect of two opposite forces. An increase in lagged employment implies that the wage must be lower to ensure employment of all previously employed persons (insiders). A larger lagged employment also implies, however, a larger current demand for labor [cf. (1)] and the wage can be higher while all previously employed persons maintain their jobs. In the simple model the former effect dominates, but in general the theoretical relationship is ambiguous [cf. the discussion by Sanfey (1995)].

Numerous studies have included lagged employment in wage equations to test this property [see Sanfey (1995) for references]. One problem with this procedure is that lagged employment often is included in models that in other respects are not consistent with the persistency hypothesis. Hence, it is not always quite clear what is tested by the procedure of plugging lagged employment into wage equations.

More importantly, this modeling strategy does not tell us much about how shocks are transmitted into wages and employment, which is the essence of the persistency hypothesis and crucial for inferring any policy conclusions from such studies.

2.3. Anticipated and Unanticipated Shocks

Considering the reduced-form wage and employment equations, we find that anticipated changes in variables relevant for the demand for labor are reflected in wages and not in employment and oppositely for unanticipated changes. Consequently, it is natural to design a test that explicitly allows us to address the question of how changes in state variables are transmitted into wages and employment.

To this end, let us formulate a slightly generalized version of the simple persistency model outlined above as

$$\begin{aligned} l_t &= \gamma_0 + \gamma_1 l_{t-1} + \gamma_2 E(\mathbf{z}_t \mid I_{t-1}) + \gamma_3[\mathbf{z}_t - E(\mathbf{z}_t \mid I_{t-1})], \\ w_t &= \beta_0 + \beta_1 l_{t-1} + \beta_2 E(\mathbf{z}_t \mid I_{t-1}) + \beta_3[\mathbf{z}_t - E(z_t \mid I_{t-1})]. \end{aligned} \tag{5}$$

The prototype insider-outsider model (3) and (4) arises as the special case in which

$$\gamma_2 = \beta_3 = 0, \qquad \gamma_1 = 1.$$

The important qualitative aspect of the basic persistency model is that anticipated changes affect wages more than employment, and oppositely for the unanticipated

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changes in the state variables. These mechanisms generalize beyond the simple model (3) and (4). If so, we maintain the central conclusion that potential improvements in employment are primarily taken out as wage increases rather than as employment increases. In particular, if $\beta_2 \neq \beta_3$, we find support that the contractual structure matters, and (i) if $\beta_2 > \beta_3$, the wage response is larger to anticipated shocks than to unanticipated ones; (ii) if $\beta_2 > 0$, wage adjustment works so as to stabilize employment ($\gamma_2 < \alpha_2$); and (iii) testing for $\beta_1 \neq 0$; $\beta_2 \neq 0$ is also a test for real-wage rigidity [cf. Blanchard and Fischer (1989)].

The empirical implementation of the tests outlined above proceeds by first considering the stationarity properties of employment as well as labor supply and employment. This is straightforward and can be done without further specification of the model. So can the efficient-market property, but because this test comes out automatically when formulating our employment model, this test is postponed. To test the effects of anticipated and unanticipated changes in state variables relevant for the demand for labor, we need to identify the state variables. To do so, we formulate an employment model in Section 4. This provides us with the information needed for the formulation of a model in Section 5 that allows us to evaluate the importance of anticipated and unanticipated shocks for wages and employment.

3. STATIONARITY PROPERTIES OF EMPLOYMENT, LABOR SUPPLY, AND UNEMPLOYMENT

The basic employment persistency model implies, cf. (4), that employment follows a random walk provided that the expectation error in predicting the state variable z is white noise.

A necessary condition for (4) to be a random walk is that employment contains a unit root, i.e., is integrated at the long-run frequency. A series that is integrated of order one at the zero or long-run frequency is a series with a stationary first difference. Such a series contains a stochastic trend, that is, a sum of all previous errors to the model, and the series is said to exhibit persistency because a shock to the series will have an everlasting influence. The commonly used test for the order of integration is the Dickey-Fuller test [see Fuller (1976)] and Table 1 reports the results of testing for first-order integration of the Danish labor force (LS), the number of employed (LD), the number of unemployed (LU), and the unemployment share LU/LS. The sample applied is $T = 44$ yearly observations form 1948 to 1991.

The results in Table 1 indicate that we cannot reject the hypothesis that these series are integrated of order one. In particular, we cannot reject the fact that employment contains a unit root at the zero frequency, as indicated by (4).

Integration of order one in LD and LU or LU/LS seems to confirm the basic prediction of the employment persistency hypothesis. However, the fact that LS and LD are integrated of order one is also consistent with the predictions of, e.g., a competitive labor market model, but this model predicts that unemployment, i.e.,

Table 1. Dickey-Fuller test 1948–1991 ($T = 44$)

Series	Dickey-Fuller[a] 't'	Augmentation (values of $j \neq 0$)
Labor force	−2.25	1
Employed	−3.30	1
Unemployed	−1.31	0
Unemployment rate	−1.26	0

[a]The auxiliary Dickey-Fuller regression is $\Delta y_t = \pi y_{t-1} + \sum_{j=1}^{k} b_j \Delta y_{t-j} + c_0 + c_1 t + \epsilon_t$ where y_t is the observed series and the augmentation is chosen so that it uses the smallest number of lag coefficients needed to make the residuals white noise. A constant and a trend were added to all regressions. The test of the null of a unit root ($\pi = 0$) against the stationary alternative ($\pi < 0$) is based on the t value corresponding to π. The distribution of the t value is nonstandard, and simulated critical values are given by Fuller (1976).

the linear combination $\mathrm{LU} = \mathrm{LS} - \mathrm{LD}$, is integrated of order zero, implying that LS and LD are cointegrated [see Engle and Granger (1987)].

That LU seems to be integrated of order one, thereby implying that LS and LD are not cointegrated, gives a model consistent with persistent unemployment, such as the insider-outsider model, a lead compared to a competitive labor market model. The power of the unit root test, however, is known to be low, and before firm conclusions are drawn, further evidence should be considered, especially evidence, which is not based only on a univariate analysis.

The preceding test has proceeded from the fact that it is a necessary condition that employment be integrated of order one for employment to follow a random walk. However, it is not a sufficient condition because it only indicates that the change in employment is stationary, whereas (4) implies that it should be white noise. The necessity of augmenting the Dickey-Fuller regression to obtain white-noise errors is, of course, an indication that employment does not follow a random walk.

As noted, the assumption that the prediction error $\mathbf{z}_t - E(\mathbf{z}_t \mid I_{t-1})$ is white noise also can be stated as the condition that the prediction error is an innovation with respect to an information set containing past prediction errors. However, this property depends crucially on the timing of the information signals belonging to the information set I_{t-1}. For instance, if the information set contains variables from year $t-2$ only, $\Delta\mathrm{LD}_t$ may be an MA(1) process $\epsilon_t + \theta\epsilon_{t-1}$, and employment changes thus will be correlated.

However, because the autocovariance of a qth-order moving-average process is zero for lags higher than q, i.e., the correlogram has a cutoff point at $q+1$, the correlogram of the first differences of the employment series of up to nine lags were computed. The estimated correlogram have the values 0.27, −0.11, −0.18, −0.33, −0.27, −0.16, 0.08, 0.28, 0.22 for lags 1, 2, 3. . . , 9, respectively. Hence, the correlogram has large values for lags 4, 5, and 8 years.

Furthermore, (4) has the stronger implication that the prediction error is an innovation with respect to whatever is in the information set I_{t-1}. We return to an explicit test of this property when an employment model has been formulated in the next section.

4. EMPLOYMENT MODEL

The next step in the testing procedure is to consider how anticipated and unanticipated changes in the state variables are reflected in wages and employment. To this end, we need to identify the state variables, and we do this by first estimating a parsimonious employment equation of the form (1). As a byproduct, this also provides an easy way to test the efficient-market property of employment changes (cf. Section 2).

4.1. Data

The data set used relates to the manufacturing sector in Denmark with a sample running from 1974:1 to 1991:4. The series includes (all variables in logs) manufacturing employment (l), labor productivity (lp), product wage (w), real raw-material price (r) as a proxy for the cost of other inputs, and exports (ex) and government expenditures (g) as proxies for international and domestic demand, respectively. The demand components are all in real terms. In the final analysis, exports and government expenditures had very similar effects, and therefore they were added together to provide a composite indicator for the demand pressure; this aggregate is denoted d. In addition, information on the unemployment compensation and the tax burden were added to the analysis, but neither variable seemed to have any significant effect on the wage-employment determination.[6]

A quick glance at plots of the series and transformations thereof, such as the first and fourth differences, of the four single quarters, and of transformations extracting the long run, the semiannual, and the annual seasonal features, indicate the existence of stochastic trends, and a somewhat varying seasonal pattern[7] in some of the series. As such, a pattern can be subjected to a formal test by testing for unit roots at long-run and seasonal frequencies, we apply the test of Hylleberg et al. (1990). The results of the HEGY test are presented in Table 2.

Although the unit roots at the seasonal frequencies all are rejected, integration of order one at the zero frequency cannot be rejected for l, ex, g, d, and r. For lp, integration of order one is barely rejected with a trend in the auxiliary regression.

We thus have found that the state variables are highly persistent, and this points to the danger of simply using the stationarity properties of employment as a test of the employment persistency hypothesis. As noted above, it is no surprise to find persistency in employment if the state variables also are persistent. The interesting question is the adjustment mechanism to changes in the state variables.

TABLE 2. HEGY test quarterly data 1974:1 to 1991:4, $T = 72$[a]

Series	t_{π_1} (long run)	t_{π_2} (semiannual)	$F_{\pi_3 \cap \pi_4}$ (annual)	Augmentation (values of $j \neq 0$)
l	−2.20	−4.86*	41.82*	0
d	−0.52	−4.28*	−47.13*	0
lp	−3.73*	−4.06*	20.06*	4
w	−2.35*	−5.25*	12.28*	2
r	−1.66	−6.92*	17.32*	0

[a]The auxiliary regression is $\Delta_4 y_t = \pi_1 y_{1t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1} + \sum_{j=1}^{k} b_j \Delta_4 y_{t-j} + c_0 + c_1 Q_{1t} + c_2 Q_{2t} + c_3 Q_{3t} + c_4 t + \epsilon_t$, where $y_{1t} = (1 + B + B^2 + B^3) y_t$ includes the long-run unit root, $y_{2t} = -(1 - B + B^2 - B^3) y_t$ includes the semiannual unit root, and $y_{3t} = -(1 - B^2) y_t$ includes the annual unit root if they exist. The test of the null of a unit root at the zero frequency ($\pi_1 = 0$) against the stationary alternative ($\pi_1 < 0$) uses the t value on π_1, whereas the test of the null of a unit root at the semiannual frequency ($\pi_2 = 0$) against the stationary alternative ($\pi_2 < 0$) uses the t value on π_2. The distributions are as the Dickey-Fuller t and the critical values are supplied by Fuller (1976) and Hylleberg et al. (1990). The latter also present the critical values for F value on the test of $\pi_3 \cap \pi_4 = 0$, which is the test for a unit root at the annual frequency. A constant, three seasonal dummies Q_{it}, $i = 1, 2, 3$, and a trend t generally were included, but the results of the case in which only a constant and the seasonal dummies were included is given in brackets. The augmentation is chosen following a procedure similar to that in Table 1. A star indicates rejection of the null of unit root against the stationary alternative at a 5% level. The 5% critical values are −3.6, −3.0, and 6.57 for the t_{π_1}, t_{π_2}, and $F_{\pi_2 \wedge \pi_4}$ tests, respectively.

4.2. Error Correction Model

Based on the result that the series are nonstationary and probably integrated of order one at the zero frequency, an error correction model for l_t was specified. The Johansen (1988, 1991), procedure, together with the two-step procedure of Engle and Granger (1987), was applied.

The Johansen maximum likelihood procedure starts by estimating the so-called interim multiplier representation [see Hylleberg and Mizon (1989)]:

$$\Delta \boldsymbol{x}_t = \Gamma_1 \Delta \boldsymbol{x}_{t-1} + \Gamma_2 \Delta \boldsymbol{x}_{t-2} + \cdots + \Gamma_{k-1} \Delta \boldsymbol{x}_{t-k+1} - \prod \boldsymbol{x}_{t-k} + \varepsilon_t, \quad t = 1, 2 \ldots T, \quad (\mathbf{6})$$

where $\boldsymbol{x}_t$ is a $p \times 1$ vector of variables observed in period t, and $\boldsymbol{\epsilon}_t$ is a $p \times 1$ vector with multivariate distribution $N(0, \Omega)$. The number of cointegrating vectors, also called the cointegrating rank (r), is the rank of $\Pi = \alpha\beta'$, where α and β are $p \times r$ matrices; α is the matrix of loadings and β contains the cointegrating vectors. Of course, neither the loadings nor the cointegrating vectors can be estimated because of lack of identification, but we can estimate the space spanned by the rows of α and the rows of β. A likelihood ratio test of the hypothesis $H_0 : \Pi = \alpha\beta'$ is applied, and the critical values from the nonstandard distribution are found in Osterwald-Lenum (1992).

As with the univariate counterpart, the Dickey-Fuller test, it is quite important when performing an ML test for the number of cointegrating vectors that k is chosen so that the errors are multivariate white noise. We found $k = 5$ to be a reasonable choice.

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TABLE 3. Likelihood ratio test for the number of cointegrating vectors[a]

	Maximum eigenvalue $-T\ \ln(1-\lambda_{r+1})$		Trace $-T\ \sum_{r+1}^{5} \ln(1-\lambda_i)$	
H_0	H_1	Test[b]	H_1	Test[b]
$r=0$	$r=1$	45.21**	$r=5$	109.5**
$r\leq 1$	$r=2$	33.18**	$r=5$	64.32**
$r\leq 2$	$r=3$	15.93	$r=5$	31.14*
$r\leq 3$	$r=4$	12.25	$r=5$	15.21
$r\leq 4$	$r=5$	2.96	$r=5$	2.9

[a]The vector autoregression model had five lags and an intercept and three seasonal dummies. The test allowed for an intercept (and seasonal dummies) in both the levels equation and the differenced equation. Hence, Table 1 of Osterwald-Lenum (1992) was used.
[b]Two asterisks indicate significance at the 5% level; one star indicates significance at the 10% level.

The test results for a system with the five variables l_t, lp_t, w_t, r_t, and d_t are presented in Table 3. At the 5% level, both the maximum eigenvalue test and the trace test point to two cointegrating vectors.

In the present setting it is natural to think of these two cointegrating vectors in terms of an employment equation and a wage equation. The result of a search based on this idea gave the following two cointegration relations:

$$l_t = 6.168 - 0.011Q_{1t} + 0.012Q_{2t} + 0.014Q_{3t} + 0.055d_t - 0.462w_t - 0.194r_t + v_{lt}, \tag{7}$$

$$R^2 = 0.61,\ \mathrm{DW} = 0.24,\ \mathrm{ADF}\colon t_\pi = -5.20\ \{\text{no intercept, augmentation } 1, 2, 3, 4, 7, 8, 12\},$$[8]

and

$$w_t = 1.218 + 0.038Q_{1t} + 0.044Q_{2t} - 0.015Q_{3t} + 0.583lp_t - 0.135r_t + v_{wt}. \tag{8}$$

$$R^2 = 0.87,\ \mathrm{DW} = 0.42,\ \mathrm{ADF}\colon t_\pi = -3.17\ \{\text{no intercept, augmentation} = 6, 7, 14, 15\},$$

where the the ADF test is the t value obtained from auxiliary Dickey-Fuller regressions of the first difference of the residuals on the lagged value of the residual with a proper augmentation. In addition, using the critical values supplied by Engle and Yoo (1987), the two cointegrating regressions cannot be rejected. The residuals from the cointegrating regresions (7) and (8) are depicted in Figures 1 and 2.

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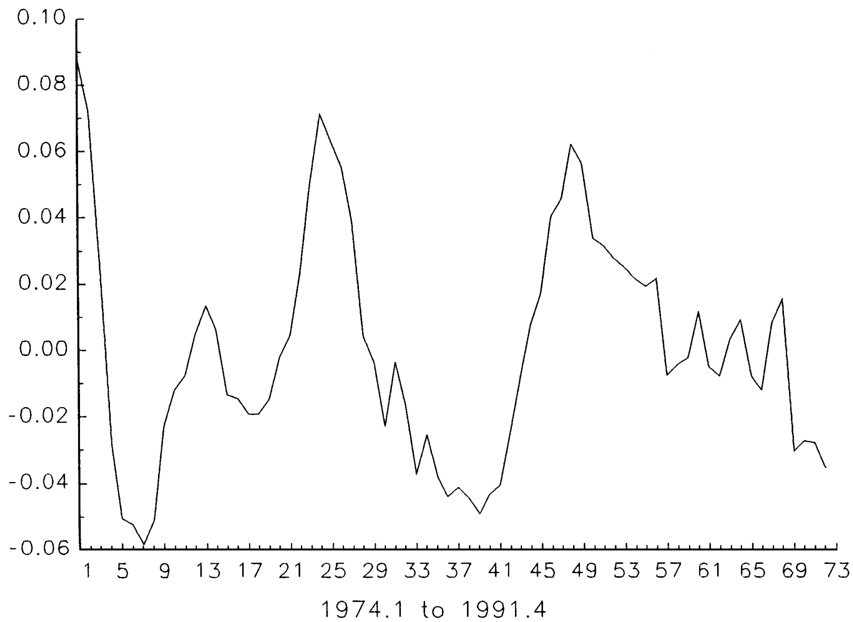

Figure 1. Residuals from cointegrating regression for employment.

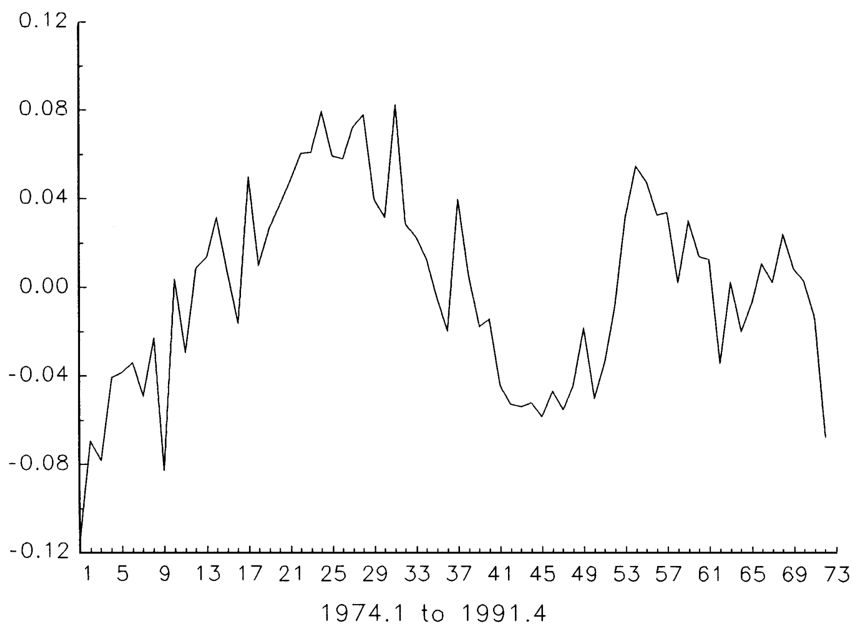

Figure 2. Residuals from cointegrating regression for wages.

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4.3. Efficient-Market Property of Employment Changes

To test the efficient-market property of employment changes, the first difference of l_t was regressed on lagged first differences of the variables in the cointegrating regression, lagged values of itself, the cointegrating errors from (7) and (8) lagged one quarter, a constant, and three seasonal dummies. The parsimonious employment equation becomes

$$\begin{aligned}\Delta l_t = {} & \underset{(0.111)}{0.333}\Delta l_{t-1} + \underset{(0.105)}{0.261}\Delta l_{t-4} - \underset{(0.105)}{0.221}\Delta l_{t-5} - \underset{(0.092)}{0.205}\Delta w_{t-1} + \underset{(0.029)}{0.068}\Delta lp_{t-2} \\ & - \underset{(0.053)}{0.083}\, v_{1,t-1} - \underset{(0.037)}{0.078} v_{w,t-1} \\ & - \underset{(0.003)}{0.015} + \underset{(0.005)}{0.004} Q_{1t} + \underset{(0.006)}{0.029} Q_{2t} + \underset{(0.006)}{0.027} Q_{3t} + e_t\end{aligned} \tag{9}$$

$R^2 = 0.78$, SE $= 0.010$, $F\{10, 55\} = 18.99[0.00]$[9]
Information criteria: SC $= -8.79$
Normality: $\chi^2(2) = 0.73$
Autocorrelation: AR 1-1: $F\{1, 54\} = 0.27[0.60]$
AR 1-5: $F\{1,54\} = 0.35[0.88]$
Autoregressive conditional heteroskedasticity: ARCH 4: $F\{4, 47\} = 0.33[0.86]$
Heteroskedasticity: X_i^2: $F\{17, 37\} = 0.64[0.84]$
Functional form: RESET: $F\{1, 54\} = 0.66[0.42]$
Length of roots in lag polynomium: Δl_t: one real and two complex: NORM $= 1.58$ and 1.69
Within-sample fit, measured in the levels of the variables, is $R^2 = 0.96$

The estimated model meets all of the usual criteria of normality, no autocorrelation, no ARCH, homoskedasticity, and functional form and the roots of the lag polynomial of the dependent first-differenced employment are all outside the unit circle as shown above. The error correction terms are both negative, and although $v_{l,t-1}$ is barely significant, it cannot be left out without causing autocorrelation problems. In case $v_{w,t-1}$ is left out, $v_{l,t-1}$ becomes strongly significant with a coefficient of -0.15.

It is seen that the estimated equation implies that the change in employment is not an innovation with respect to an information set that includes lagged changes of employment, wages, and productivity, and the lagged cointegrating relations. Hence, because these variables would be in any sensible information set, the efficient-market property of (4) is not congruent with the data. With some justification, it could be argued here that a quarter may be too short a period for the information to arrive, but the lags in (9) are actually quite long.

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Notice that the estimated equation (9) suggests that factors such as export and government expenditures, and raw material prices have no short-run effects on employment, but only long-run effects through the cointegrating relations. The long-run effects on employment of an increase in demand are positive, whereas the long-run effect of increases in wages or raw-material prices is negative. In the short run, wage increases depress employment, whereas productivity increases have a positive effect. Employment will fall if the wage is above its long-run sustainable value seen in relation to productivity and raw-material prices [see (8)] or if employment is above its long-run equilibrium value according to (7).

5. ADJUSTMENTS TO ANTICIPATED AND UNANTICIPATED SHOCKS

Having identified the set of state variables in the preceding section, we now can proceed to test how anticipated and unanticipated changes in these state variables affect wages and employment so as to test the basic implications of the insider model for the adjustment process.

The model we use for this test is a version of (5) that allows for more lags in past employment to obtain a better representation of the dynamics. Notice that this model also allows us to test the prediction that lagged employment exerts a negative influence on wages.

In setting up the wage-employment model, we assume that the first differences of d_t, r_t, and lp_t are strongly exogenous for the parameters of interest of the model, and we separate the observed exogenous state variables into an expected part and an unexpected part by estimating a multivariate vector autoregression for $(\Delta d_t, \Delta r_t, \Delta lp_t)'$. In addition, we assume that the long-run relations are as determined in (7) and (8). See Wickens (1982), Bean (1986), and Pagan (1986) for discussions of cases in which these assumptions are not met.

The predictions of the VAR model for the first difference of the state variables, which turned out to be of fifth order, then are used as the anticipated values and the prediction errors as the unanticipated values,[10] in a multivariate error-correction model for $(\Delta l_t, \Delta w_t)'$, with the lagged error-correction terms $v_{l,t-1}$ and $v_{w,t-1}$ in both equations. Table 4 shows the results of a general-to-specific modeling exercise allowing lagged and unlagged values of both the anticipated and the unanticipated exogenous variables, and lagged and unlagged endogenous variables by using both ordinary least squares (OLS) and full-information maximum likelihood (FIML).

The approach here is similar to that used in testing whether anticipated nominal demand changes have real effects [see, e.g., Mishkin (1983)]. The only important difference is that we have to allow for the state variable to be multidimensional whereas tests of New-Classical models usually proceed with a one-dimensional state variable.

The OLS estimates and the FIML estimates are almost identical because the simultaneity of the resulting wage-employment model only comes through the disturbance covariance matrix, and the covariance is small. More important, this implies that the equations in Table 4 are the restricted reduced forms and they can be interpreted as such.

TABLE 4. Estimated wage-employment model[a]

	OLS		FIML	
Variable	Wage Δw_t	Employment Δl_t	Wage Δw_t	Employment Δl_t
Δl_{t-1}		0.329		0.371
		(0.118)		(0.103)
Δl_{t-2}	−0.248		−0.270	
	(0.109)		(0.097)	
Δw_{t-1}		−0.149		−0.190
		(0.100)		(0.088)
Δw_{t-5}	0.231		0.241	
	(0.089)		(0.079)	
$v_{l,t-1}$	0.074	−0.076	−0.077	−0.049
	(0.042)	(0.045)	(0.038)	(0.040)
$v_{w,t-1}$	−0.112	−0.104	−0.117	−0.097
	(0.040)	(0.038)	(0.036)	(0.033)
Δlpu_{t-2}		0.113		0.120
		(0.046)		(0.040)
Δdu_t		0.143		0.166
		(0.082)		(0.071)
Δdu_{t-3}	0.161		0.163	
	(0.078)		(0.069)	
Δru_t	−0.108		−0.106	
	(0.035)		(0.032)	
Δda_{t-3}	−0.239		−0.207	
	(0.104)		(0.093)	
Δra_t	−0.123		−0.126	
	(0.041)		(0.036)	
R^2	0.76	0.76	Ln $\|\Omega\|$[l]	−19.19
SE	0.01	0.01	Likelihood $-c$	14,655.96
F value	14.96[b] [0.00]	19.14[g] [0.00]	Ln $\|Y'Y/T\|$	−17.66
AR 1–4 F		0.64[h] [0.64]	VAC	0.15
Normality χ_2^2	0.90[c] [0.47]	0.52	Trace correlation	0.78
Hetoskeda-sticity X_i^2 F	1.03	0.69[i] [0.78]	LR test for overidentifying	
ARCH 1–4 F	0.42[d] [0.97]	0.18[j] [0.26]	restrictions.	
RESET F	0.80[e] [0.53]	2.06[k] [0.16]		
	0.15[f] [0.70]		χ_{16}^2	22.19

[a] An intercept and three seasonal dummies are included in all regressions. The unanticipated values or prediction errors are indicated by u after the variable name; a indicates an anticipated value. Numbers in parentheses are standard deviations; number in brackets are the p values, i.e., the probabilities of getting a larger value than the test statistics. [b] Degrees of freedom (DOF) $= 11, 51$. [c] DOF $= 4, 47$. [d] DOF $= 19, 31$. [e] DOF $= 4, 43$. [f] DOF $= 1, 52$. [g] DOF $= 9, 54$. [h] DOF $= 4, 50$. [i] DOF $= 15, 38$. [j] DOF $= 4, 46$. [k] DOF $= 1, 53$. [l] Ω is the variance covariance matrix of the reduced-form residuals.

https://doi.org/10.1017/S1365100598009043 Published online by Cambridge University Press

Included in the resulting estimated employment equation is now the unanticipated increase in productivity (with a lag of two quarters) and the unanticipated change in demand while none of the anticipated values appear. An F-test for adding the predicted values of the quarterly change in productivity, demand, and raw-material prices gives $F_{3,51} = 0.74[0.53]$, indicating that the anticipated shocks in the state variables should not be in the employment equation. Similarly, testing for the inclusion of up to three lags of the unanticipated changes in the raw-material prices gives $F_{4,49} = 0.21[0.93]$, implying that these variables should not be included either. The two error-correction terms both have negative coefficients, and the employment equation is only slightly different from the one estimated earlier, where no distinction was made between anticipated and unanticipated changes in the state variables. Hence, the employment equation is in accordance with the predictions of the persistency hypothesis.

However, the resulting estimated wage equation contains both anticipated and unanticipated values of demand and raw-material prices. In fact, the division of the raw material prices into an anticipated part and an unanticipated part has no effect at all, because the coefficients to the two parts are not significantly different. A test of this is performed by including the observed series r_t in the regression instead of ru_t or ra_t and computing an LM test for the inclusion of either one of the two variables in the regression. F values of $F\{_{1,51}\} = 0.087$ for both of these indicate that we cannot reject the hypothesis of a common coefficient. This, of course, contradicts the persistency model. Both anticipated and unanticipated exogenous demand changes affect wages, and whereas the anticipated component somewhat surprisingly affects wages negatively, the unanticipated component has the expected positive effects on wages as well as on employment. This implies that only the incumbent workforce is affected by expected changes in demand, whereas both employment and wages are affected by the unanticipated shocks to demand. In addition, the wage decreases if it is above its long-run equilibrium and increases if employment exceeds its long-run equilibrium. This implies that the change in wage will be negative in a case in which the wages in the preceding period are too high, and the employment too low compared to their long-run equilibrium values. The dynamic adjustment mechanisms of wages and employment thus are qualitatively different.

Thus, the prediction of the extreme version of the persistency model that only anticipated changes affect wages and only unanticipated changes affect employment cannot be supported by the data, but a slightly less extreme version seems to be congruent with the data. Whereas only unanticipated shocks to the state variables enter the employment equation, both anticipated and unanticipated shocks enters the wage equation.

Lagged employment enters the wage equation as well, partly through the error-correction term, where the effect arises if employment is different from the long-run value, and partly directly. In addition, the coefficient is negative as predicted by the insider-outsider theory.

Previous studies for Denmark have reached different conclusions on this issue. In a study of wage and employment determination in the manufacturing sector,

https://doi.org/10.1017/S1365100598009043 Published online by Cambridge University Press

Andersen and Risager (1991) found no effects of lagged employment, but when Risager (1992) studied insider-outsider influences for the wage setting of skilled and unskilled men, he found a negative effect of lagged employment for the latter group. However, it is, surprising that the effect is found for unskilled but not for skilled men, as also noted by Risager.

The evidence reported in Table 4 indicates that wages also adjust to unexpected changes in the state variables; hence, there is a possibility of adjusting the wage within the period (wage drift) despite wage negotiations at regular intervals. This raises the question whether the adjustment within the period is so flexible that the contractual structure does not matter for the final wage settlement. To evaluate this hypothesis, we could estimate the wage equation by imposing the restriction that anticipated and unanticipated changes in the state variable (here the change in the raw-material price, Δra_t and Δru_t, and the change in the demand, da_{t-3} and du_{t-3}) have the same effect on the change in the wage. This was found to be the case for the raw-material prices, but it is quite obvious from the result shown in Table 4 that the coefficients are different for the demand variable. Hence, the contractual structure seems to matter. Traditionally, this issue is addressed by running regressions to test whether centrally negotiated wage changes and drift are perfect substitutes. The present approach offers a more direct test.

6. CONCLUDING REMARKS

The present study has shown that employment and unemployment are highly persistent, but so are the state variables driving labor demand. Hence, it is unclear whether persistency is due to endogenous propagation mechanisms or the properties of the shocks. Accordingly, the policy implications of persistency are unclear.

Using the test strategy developed, a detailed study of wage and employment determination reveals that endogenous propagation mechanisms are at stake as indicated by the fact that anticipated changes in state variables of relevance for labor demand are reflected only in wages, leaving employment unaffected. This is in accordance with the basic persistency hypothesis and reveals that an endogenous propagation mechanism arises via wage formation. This conclusion is reinforced by the finding that lagged employment exerts a negative difference on current wages. Unanticipated changes in the state variables affect both wages and employment, indicating that although wages are preset, there is some scope for adjustment within the period.

The policy conclusion of these findings is that there are structural impediments, which makes it difficult to lower unemployment. Policies directed only at boosting labor demand run the risk of fueling wage increases, leaving employment unaffected if they are not accompanied by measures directed at wage formation.

NOTES

1. The sensitivity of labor supply to wages generally is found to be very weak or absent, at least in the short run.

2. It is well-known that the insider-outsider model in general does not imply employment persistency. See Sanfey (1995) for a recent survey of the theoretical and empirical literature on the insider-outsider model.

3. The method developed in this paper has been used by Jansson (1995) to test for persistency in Swedish employment.

4. See Blanchard and Fisher (1989), Lindbeck and Snower (1988), and Andersen and Velter (1994).

5. It is well known that this property depends crucially on the specific way in which membership is modeled [see Holmlund (1991) and Sanfey (1995)].

6. This confirms that wage determination is driven primarily by inside variables in accordance with the insider-outsider model [see, e.g., Nickell and Kong (1992)].

7. See Hylleberg (1992, 1994) for a presentation of different helpful graphical tools.

8. Autocorrelation still exists at lag 10. But this could not be removed by any feasible augmentation. The result therefore must be interpreted with care.

9. The numbers in { } are degrees of freedom; the numbers in brackets are *p* values, i.e., the probabilities of getting a larger value than the value of the test statistics.

10. Although the individual errors in the three variable VAR's are contemporaneously correlated, they are innovations with respect to the information set applied. We treat the errors from each equation as the unanticipated shock to the dependent variable, but test the common effect as well.

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APPENDIX

TABLE A-1. Number of employed in manufacturing (logs)[a]

	Quarter			
Year	1	2	3	4
1974	5.765	5.759	5.724	5.652
1975	5.617	5.620	5.617	5.598
1976	5.613	5.631	5.638	5.631
1977	5.631	5.631	5.620	5.602
1978	5.595	5.617	5.624	5.613
1979	5.606	5.628	5.642	5.631
1980	5.609	5.613	5.602	5.545
1981	5.521	5.517	5.545	5.525
1982	5.505	5.529	5.525	5.501
1983	5.493	5.513	5.517	5.509
1984	5.505	5.541	5.561	5.561
1985	5.568	5.613	5.628	5.620
1986	5.620	5.635	5.635	5.609
1987	5.591	5.602	5.602	5.583
1988	5.557	5.572	5.572	5.568
1989	5.545	5.568	5.583	5.576
1990	5.553	5.572	5.583	5.561
1991	5.525	5.541	5.545	5.523

[a] The series is denoted l, and it is the logarithm of the number of employed in the Danish manufacturing sector.
Source: Danmarks Statistik, SE.

TABLE A-2. Real product wage in manufacturing (logs)[a]

	Quarter			
Year	1	2	3	4
1974	1.488	1.532	1.525	1.552
1975	1.580	1.634	1.633	1.654
1976	1.652	1.689	1.686	1.706
1977	1.698	1.720	1.721	1.734
1978	1.738	1.756	1.764	1.793
1979	1.790	1.804	1.786	1.812
1980	1.776	1.778	1.780	1.784
1981	1.770	1.763	1.756	1.769
1982	1.739	1.768	1.751	1.759
1983	1.775	1.790	1.765	1.762
1984	1.748	1.764	1.762	1.760
1985	1.741	1.759	1.771	1.795
1986	1.802	1.836	1.861	1.879
1987	1.898	1.930	1.925	1.947
1988	1.932	1.948	1.950	1.962
1989	1.936	1.933	1.935	1.939
1990	1.947	1.970	1.963	1.975
1991	1.967	2.002	2.000	2.007

[a] The series is denoted w, and it is the logarithm of the average hourly manufacturing wage divided by the price of manufacturing output.
Source: Danmarks Statistik.

TABLE A-3. Real raw-material price (logs)[a]

	Quarter			
Year	1	2	3	4
1974	0.2370	0.2195	0.2011	0.1646
1975	0.1079	0.0832	0.0871	0.1056
1976	0.1290	0.1369	0.1452	0.1365
1977	0.1526	0.1778	0.1526	0.1463
1978	0.0957	0.0735	0.0560	0.0417
1979	0.0716	0.1580	0.2865	0.3343
1980	0.4495	0.5056	0.4940	0.5313
1981	0.6108	0.6787	0.6735	0.6134
1982	0.6387	0.6331	0.6427	0.6533
1983	0.6211	0.5882	0.6165	0.6337
1984	0.6490	0.6318	0.6429	0.6561
1985	0.6643	0.6327	0.5700	0.5588
1986	0.4476	0.3057	0.2364	0.2334
1987	0.2195	0.1952	0.2129	0.2022
1988	0.1735	0.1912	0.2148	0.2128
1989	0.2515	0.2594	0.2441	0.2339
1990	0.1910	0.1387	0.2136	0.2758
1991	0.1792	0.1575	0.1521	0.1430

[a] The series is denoted r, and it is the logarithm of the raw-material price divided by the price of manufacturing output.
Source: Danmarks Statistik.

TABLE A-4. Labor productivity (logs)[a]

	Quarter			
Year	1	2	3	4
1974	0.6483	0.6324	0.7319	0.6799
1975	0.6466	0.7155	0.8394	0.8114
1976	0.8512	0.7569	0.9114	0.8525
1977	0.7709	0.7722	0.9104	0.9456
1978	0.7641	0.8466	0.9281	0.9306
1979	0.8499	0.8621	0.9600	0.9586
1980	0.8938	0.9014	0.9777	0.9589
1981	0.9540	0.9616	0.9615	1.036
1982	0.9361	0.9918	1.098	1.112
1983	0.9660	1.032	1.135	1.104
1984	1.069	1.097	1.198	1.168
1985	1.084	1.077	1.199	1.195
1986	1.070	1.139	1.238	1.201
1987	1.097	1.095	1.204	1.238
1988	1.141	1.216	1.277	1.299
1989	1.202	1.268	1.305	1.322
1990	1.240	1.227	1.346	1.319
1991	1.245	1.298	1.424	1.500

[a] The series is denoted lp, and it is the logarithm of the (manufacturing sale + inventory investments)/number of working hours in the Danish manufacturing sector. This construction was necessary, because data for the manufacturing production are not directly available.
Source: Danmarks Statistik.

TABLE A-5. Exogenous demand (logs)[a]

	Quarter			
Year	1	2	3	4
1974	4.597	4.654	4.694	4.729
1975	4.722	4.772	4.777	4.817
1976	4.836	4.883	4.902	4.937
1977	4.974	4.961	4.999	5.033
1978	5.037	5.091	5.107	5.132
1979	5.162	5.208	5.257	5.318
1980	5.377	5.380	5.413	5.439
1981	5.509	5.555	5.592	5.618
1982	5.664	5.695	5.712	5.743
1983	5.762	5.773	5.789	5.832
1984	5.839	5.855	5.882	5.897
1985	5.930	5.941	5.957	5.944
1986	5.908	5.943	5.914	5.920
1987	5.940	5.969	6.000	6.018
1988	6.038	6.037	6.056	6.089
1989	6.087	6.147	6.143	6.156
1990	6.163	6.175	6.180	6.208
1991	6.197	6.224	6.234	6.245

[a] The series is denoted d, and it is the logarithm of the sum of government expenditures and export.
Source: MONA Nationalbanken.

Table A-6. Labor supply[a]

Year	Quarter 1	2	3	4
1948	1996	2020	2054	2084
1952	2088	2111	2122	2130
1956	2131	2133	2135	2151
1960	2176	2208	2240	2256
1964	2282	2313	2294	2279
1968	2301	2356	2371	2403
1972	2408	2430	2462	2468
1976	2514	2563	2615	2615
1980	2626	2653	2683	2711
1984	2746	2783	2818	2842
1988	2849	2852	2845	2846

[a] The series is denoted LS, and it is the number of workers available.
Source: Danmarks Statistik.

Table A-7. Unemployed[a]

Year	Quarter 1	2	3	4
1948	62.00	72.30	67.50	75.20
1952	97.40	76.20	66.60	79.80
1956	91.20	87.00	82.10	54.20
1960	38.90	31.10	28.30	40.60
1964	23.40	19.90	22.90	27.90
1968	49.40	39.40	30.10	38.90
1972	38.70	25.40	57.70	130.6
1976	134.1	164.0	190.6	161.8
1980	183.8	243.0	262.8	283.0
1984	276.3	251.8	220.4	221.9
1988	243.9	264.9	271.7	296.1

[a] The series is denoted LU, and it is the number of unemployed per thousand people.
Source: Danmarks Statistik.