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Evaluating the Persistence and Structuralist Theories of Unemployment from a Nonlinear Perspective

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Abstract. *This paper assesses empirically two competing unemployment theories. It identifies one structural break in U.K. and German unemployment around 1980 that is more severe in both absolute and relative terms than that for the United States in 1973. This offers support for the structuralist theory. A nonlinear (TAR) model is used to capture fast-up, slow-down unemployment dynamics. Impulse response functions suggest that the half-lives of shocks are longer in the postbreak subsamples, especially in Europe, which places the persistence theory closer to the mark. We conclude that elements from both theories are needed for an adequate account of unemployment dynamics.*

Keywords. asymmetries, structural breaks, hysteresis, threshold autoregression

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1 Introduction

The sustained rise of unemployment in many of the OECD countries in the last two decades has sparked a growing literature on unemployment dynamics that has enriched our understanding of the strengths and limitations of the natural rate of unemployment theory set out by Friedman (1968) and Phelps (1968).[1] The current literature offers two competing sets of explanations for changes in medium- to long-term unemployment. One emphasizes the importance of the adjustment process toward equilibrium and, more particularly, its sluggishness or persistence. The other focuses more on changes in the underlying equilibrium unemployment rate. We refer to the former as the persistence school and to the latter as the structuralist school.

A proper understanding of high unemployment is of great importance for policymaking. If unemployment is high because of persistence, monetary policy can make a contribution without an immediate risk of inflation. Although a transient recession could, absent intervention, leave its remnants for years, equally an expansionary monetary policy could have a sustained beneficial effect. By contrast, if high unemployment is caused by autonomous movements of the natural rate itself, expansionary monetary policy will be inflationary, because it will not affect the microeconomic foundations of the underlying labor market equilibrium.

Although the structuralist school has recently gained converts—Blanchard (1999) is a prominent example—there is a lack of empirical work attempting to discriminate between the two approaches. This has left a gap in our knowledge of the causes of high unemployment. As a result, economists find themselves disagreeing on such fundamental policy issues as labor market reforms, the European single currency, and the current formulation of monetary policy by the European Central Bank. This paper seeks to fill that lacuna by combining two methodological insights in studying the time-series properties of U.K., U.S., and German unemployment.

Both of these insights entail departures from the standard linear framework traditionally used to analyze unemployment dynamics. First, we investigate whether a structural break or modeling equilibrium unemployment as piecewise linear can shed light on its apparent hysteresis in the unit root sense. In this respect the paper builds on Bianchi and Zoega (1998) and Papell, Murray, and Ghiblawi (1999), which provide evidence that neglecting infrequent shifts in mean gives an exaggerated picture of unemployment persistence. Our data suggest that the apparent hysteresis in U.K. unemployment and, more particularly, the high rates of recent decades can partly be attributed to a single regime shift in 1980, possibly prompted by changes in fundamentals. Moreover the size of the break in the European economies was more severe than in the United States in both absolute and relative terms, which places the structuralist school closer to the mark.

Second, the paper takes into account business cycle asymmetries in explaining unemployment movements. To this end, the Blanchard and Summers (1986) conceptualization of hysteresis is extended to permit nonlinear dynamics. The paper employs the momentum threshold autoregressive (M-TAR) model introduced by Enders and Granger (1998), which can capture stylized features of unemployment behavior. M-TAR estimates indicate that unemployment rises more rapidly than it falls across time and countries. Nonlinear impulse response functions suggest substantially longer half-lives in the falling unemployment regime, which leaves the persistence school closer to the mark. Moreover, shocks in the postbreak subsamples are found to be more persistent than in the earlier subsamples, particularly in the European unemployment series.

One implication is that neither theory on its own can adequately explain unemployment hysteresis. Although the sluggishness of falling unemployment, especially in recent decades, can explain short- to medium-term high unemployment, it does not provide a complete rationale for the European unemployment experience during the past two decades. A single structural break, possibly associated with changes in

[1] See Bean (1994) and Blanchard and Katz (1997).

fundamentals, can shed light on such longer-term movements but is not the whole story either. The ineluctable conclusion is that elements from both theories are required for a satisfactory account of medium- to long-term unemployment dynamics. A single structural break combined with cyclical asymmetries can account for the apparent $I(1)$-ness of unemployment in all three economies and especially in Europe.

Section 2 of the paper outlines the major differences between the persistence and structuralist theories and sets out a simple economic model that encompasses both theories. Section 3 discusses the econometric approach, and Section 4 analyzes the empirical findings and implications. Section 5 concludes.

2 Economic Theory

2.1 The persistence and structuralist schools

Persistence school contributions such as those of Lindbeck and Snower (1988), Layard, Nickell, and Jackman (1991), and Karanassou and Snower (1998) emphasize the role of dynamic adjustment in the natural rate of unemployment. We define the latter as the steady-state rate consistent with correct expectations. These studies are concerned with the propagation of transitory nominal and real shocks. Structuralist school adherents such as Pissarides (1990), Phelps (1994), Nickell and Layard (1998), and Phelps and Zoega (1998) focus on autonomous movements of the natural rate itself due to the interplay of changes in real macroeconomic variables and institutions.

Many economists belonging to the persistence school claim that labor market flexibility is the key to good unemployment performance.[2] They argue that temporary shocks such as cyclical movements in unemployment often translate into medium-term (five- to ten-year) high unemployment, making the exact value taken by the natural rate less important. Along these lines, Layard, Nickell, and Jackman (1991), Blanchard and Summers (1986), and others have described the high European unemployment problem as hysteresis measured by the sum of the coefficients of lagged unemployment in a linear AR(MA) process.

Following this definition many studies have characterized unemployment hysteresis by means of a (near) unit root in a linear process, that is, a flat-bottomed Liapunov function for the dynamic system.[3] More recently, hysteresis has been viewed as a nonlinear phenomenon in which fixed and sunk adjustment costs make current unemployment a function of the highs and lows of past labor demand (Cross, 1994, 1995). Although the latter can be reconciled with the recent time-series literature on unemployment that highlights asymmetric adjustment, this literature has not directly addressed the issue of sustained high European unemployment.[4] To fill this gap, this paper adopts the Blanchard and Summers (1986) conceptualization of hysteresis, although in the broader sense of permitting nonlinear dynamics. This more general framework facilitates models of unemployment with distinct degrees of persistence over the phases of the business cycle.

The structuralist school embraces those economists who see the natural rate and associated equilibrium paths as endogenous and affected by market forces like any other economic variable. They have derived an equilibrium theory of unemployment movements—a moving-natural-rate theory of changes in actual unemployment as in Pissarides (1990) and Phelps (1994)—while downplaying, although far from ignoring, the role of slow adjustment to this equilibrium. Such slow adjustment is caused, for instance, by the cost of hiring and training new workers or by the cost of adjusting to new capital. Labor market equilibrium is modeled by introducing real-wage rigidity with full microfoundations. Efficiency wage models (Calvo, 1979; Shapiro and

[2]See, for instance, Elmeskov (forthcoming).

[3]See, for instance, Blanchard and Summers (1986). Note, however, that the concept of hysteresis in unemployment was introduced by Phelps (1972) and defined rather differently as the dependence of the equilibrium rate on the path followed by actual unemployment towards this equilibrium.

Within the linear tradition, an alternative characterization of unemployment persistence is as a long memory process by means of models belonging to the fractionally integrated class. Recent examples include Koustas and Veloce (1996), Crato and Rothman (1996), and van Dijk, Franses, and Paap (2001).

[4]See Açemoglou and Scott (1994), Enders and Granger (1998), and Koop and Potter (1999) on dynamic asymmetries.

Stiglitz, 1984; Katz, 1986) and models of union behavior (Layard and Nickell, 1986) are two examples. In equilibrium, a downward-sloping labor demand or price-setting curve intersects an upward-sloping wage curve in the real-wage/employment plane.

Any shift in either curve then translates into a change in the rate of equilibrium unemployment. Differences in average unemployment across epochs and countries can be traced to differences in the movement and position of either curve. The labor demand curve can potentially shift because of changes in factors such as real interest rates (Phelps, 1994; Blanchard, 1999). The position of the wage curve is determined by factors such as the generosity of the unemployment benefit welfare system, other forms of nonwage income such as rent and interest, the family network, and the consumption tax wedge. According to the natural-rate hypothesis, however, the equilibrium rate is independent of the adjustment path taken by unemployment and thus of all current and past monetary variables.

The difference of opinion between the two schools of thought hinges on whether high average unemployment can be traced to a very slow speed of adjustment toward the steady state, for a given size of transitory unemployment shock, or to an autonomous movement of the steady state itself. In brief, whereas the persistence school implicitly assumes that unemployment behaves like a (near) unit root process, the structuralist school assumes that unemployment is mean-reverting toward an occasionally changing natural rate. Therefore, we might be able to assess the empirical plausibility of these competing schools in the light of the high European unemployment of recent decades by investigating whether a change from a low to a high regime, such as the one that occurred in Germany and the United Kingdom around 1980, corresponds to a fall in the speed of adjustment. Positive results on this count would support the persistence school. On the contrary, sharp differences in average unemployment across countries during long periods may be accounted for by structural breaks in the steady-state path of a stochastic process, which play a key role in the structuralist school.

2.2 A simple economic model

A simple model is outlined drawing on an endogenous natural rate of unemployment (the steady-state equilibrium rate of unemployment) in Pissarides (1990), Layard, Nickell, and Jackman (1991), and Phelps (1994). Let the latter be denoted by $u_t^* = \boldsymbol{\lambda}'\mathbf{x}_t$, where $\boldsymbol{\lambda}$ is a vector of coefficients and $\mathbf{x}_t$ is a vector of variables, the fundamentals of the natural rate. These include elements of the welfare system such as the level and duration of unemployment benefits, productivity growth, the real interest rate, wealth, and the level and variance of educational achievement.

The economy can depart from its natural rate for significant periods because of nominal shocks, transitory real shocks, and changes in fundamentals that only gradually affect actual employment. The following error correction equation parsimoniously summarizes these dynamic interactions:

$$\Delta u_t = \alpha_0 + \alpha_1 \Delta\pi + \rho[u_{t-1}^* - u_{t-1}] + \varepsilon_t, \;\; \alpha_0 > 0, \alpha_1 < 0, \rho > 0 \qquad (2.1)$$

where u_t denotes the rate of unemployment, π denotes inflation, and ε_t is a disturbance term. Changes in u_t can be traced either to nominal demand shocks proxied by $\Delta\pi$ and/or other nonmonetary transitory shocks represented by ε_t or to a catch-up effect. The latter pertains when unemployment moves at speed ρ toward a (possibly changed) steady state, the new natural rate of unemployment.

Equation (2.1) implies that medium-term increases in the unemployment rate can be caused either by an adverse demand shock ($\Delta\pi < 0, \varepsilon_t > 0$) combined with a very low value of ρ (high persistence or hysteresis), or by autonomous increases in the equilibrium rate due to changes in some of its (nonmonetary) fundamentals, or by both. In explaining differences in average unemployment over decades, we would expect the high-unemployment economies to have a low value of ρ when compared to low-unemployment economies, assuming that they suffered similarly sized demand shocks in the past. Alternatively, they might have different values in the components of $\mathbf{x}_t$, such as a lower rate of productivity growth that shifts the equilibrium rate. The purpose of this paper is to assess empirically the relative importance of these two possibilities.

It seems plausible that the value of ρ may depend on the nature, upward or downward, of the "momentum" or inertia built into unemployment over time, and this provides the motivation for the momentum threshold autoregression framework. The latter seeks to capture the asymmetric behavior of unemployment over the business cycle, which is characterized by rapid increases during recessions and slower declines during expansions. This dynamic asymmetry can be explained by asymmetric labor adjustment costs and insider-outsider relationships. For instance, Lindbeck and Snower (1987) suggest it may be caused by asymmetric union membership rules. It takes more time to gain union membership, once hired, than to lose it following an involuntary dismissal. An unexpected fall in labor demand reduces union membership almost instantaneously, and the union leadership is then driven to protect the employment of the remaining members. By contrast, a positive labor demand shock does not raise union membership as quickly, and the union leadership will only gradually seek to protect the employment and interests of the newcomers.

3 The Econometric Framework

One of the themes of this paper is that neglected regime shifts and asymmetries may give an exaggerated impression of unemployment hysteresis. The former is motivated by Granger and Teräsvirta (1999) and Granger and Hyung (1999), who demonstrate that some properties similar to those of a long memory process may be an artifact of occasional structural breaks in a linear process. In the context of unemployment this is exemplified, for instance, by Bianchi and Zoega (1998), who employ a Hamilton Markov switching model to identify occasional shifts in the mean rate of 15 OECD economies in 1970–1996. They show that once these shifts are accounted for, the remaining persistence is significantly smaller, especially in the EU economies. Similarly, Papell, Murray, and Ghiblawi (1999) provide evidence for 16 OECD countries that the persistence of unemployment falls dramatically once structural change is incorporated. Both of these studies interpret their evidence as support for the natural-rate hypothesis. One important element that is missing in these contributions, however, is the stylized steep increases in unemployment during recessions and slower declines during expansions. This paper seeks to complement the regime shift literature with business cycle asymmetries.

3.1 A single structural break

We deploy the Zivot and Andrews [ZA] (1992) sequential unit root testing procedure, which allows for one break (at an unknown point T_B) under the stationarity alternative hypothesis. The estimated break point $\hat{T}_B$ is the date that gives the least favorable result for the unit root null. The most general model employed by ZA is

$$\Delta u_t = \theta_0 + \theta_1 t + \theta_2 DU_t + \theta_3 DT_t + \rho u_{t-1} + \sum_{i=1}^{k} \beta_i \Delta u_{t-i} + \varepsilon_t \tag{3.1}$$

for $t = 1, \ldots, T$, where $\varepsilon_t \sim$i.i.d.$(0, \sigma^2)$, and

$$DU_t = \begin{cases} 1 & \text{if } t > T_B \\ 0 & \text{otherwise} \end{cases} \qquad DT_t = \begin{cases} t - T_B & \text{if } t > T_B \\ 0 & \text{otherwise} \end{cases}$$

are level and trend dummies, respectively. Two alternative models include either DU_t or DT_t only. Sequential t-statistics for ρ, denoted $\tau(\gamma)$, with $\gamma = T_B/T$, are obtained for $k + 2 \leq T_B \leq T - 1$, and the break date estimate is given by $\hat{T}_B = \arg\min_\gamma \tau(\gamma)$.

3.2 An empirical M-TAR model

Our empirical methodology seeks to combine a single regime shift in the equilibrium level and asymmetries in the speed of adjustment. To this end, the observed series is divided into subsamples at the identified break point, and since standard linear Gaussian parameterizations cannot represent asymmetries, a nonlinear model is fitted on each of the subsamples. Those nonlinear models most frequently applied to explore unemployment are state-dependent representations either of the threshold autoregressive (TAR) type developed by Tong (1978) and Tong and Lim (1980) or the smooth-transition autoregressive (STAR) variety developed by Granger and Teräsvirta (1993). This paper puts forward a TAR model, variants of which have been applied to U.S. unemployment rates by Caner and Hansen (1998), Hansen (1997), Koop and Potter (1999), and Rothman (1998), among others.[5]

The determination of the natural rate of unemployment u_t^* in (2.1) is not modeled. Instead, once the regime shift is accounted for by splitting the series into two subperiods, u_t^* is proxied separately for each on the assumption that it is either a constant or changes smoothly over time in response to changes in its fundamentals. To formulate the nonlinear model, let $q_t(\mathbf{w}, d)$ represent a weighted sum of past unemployment changes:

$$q_t(\mathbf{w}, d) = w_1 \Delta u_{t-1} + \cdots + w_d \Delta u_{t-d} \tag{3.2}$$

where $\mathbf{w}' = (w_1, \ldots, w_d) \geq 0$ are the weights and $d \geq 1$ is a (finite-integer) threshold delay parameter. Consider the following univariate error correction process whose adjustment speed is permitted to take two values depending on the state (either nondecreasing or decreasing) of $q_t(\mathbf{w}, d)$:

$$\Delta u_t = I_t \rho_1 (u_{t-1} - \mu) + (1 - I_t)\rho_2 (u_{t-1} - \mu) + \sum_{i=1}^{k} \beta_i \Delta u_{t-i} + \varepsilon_t \tag{3.3}$$

where $\varepsilon_t \sim$ i.i.d.$(0, \sigma^2)$ and I_t is the indicator function:

$$I_t = \begin{cases} 1 & \text{if } q_t(\mathbf{w}, d) \geq 0 \text{ (RU regime)} \\ 0 & \text{if } q_t(\mathbf{w}, d) < 0 \text{ (FU regime)} \end{cases}$$

If $-2 < (\rho_1, \rho_2) < 0$, the series u_t is covariance-stationary and μ plays the role of attractor. The estimated μ is our proxy for u_t^*. Within this framework $q_t(\mathbf{w}, d)$ is the threshold or switching variable. The strength of the attraction is thus measured by ρ_1 in a rising unemployment (RU) regime and ρ_2 in a falling unemployment (FU) regime. Model (3.3) with $d = 1$ and $w_1 = 1$ was introduced by Enders and Granger (1998) and called a momentum threshold autoregression (M-TAR).

Varying the weighting scheme in (3.2) results in different measures of past unemployment growth. For instance, for $w_1 = w_2 = \cdots = w_{d-1} = 1$ and $w_d = 0$, $q_t(\mathbf{w}, d)$ collapses to a long difference (LD)

$$q_t(LD, d) = \Delta u_{t-1} + \cdots + \Delta u_{t-(d-1)} \equiv u_{t-1} - u_{t-d}$$

which represents cumulative past changes. An LD with large d was proposed by Caner and Hansen (1998) in the context of unemployment as a less noisy alternative than a simple (lagged) first-difference or $q_t(LD, 2) = \Delta u_{t-1}$.[6]

Two new threshold variables consisting of convex combinations of past unemployment changes are proposed. In one, recent changes are more relevant in triggering the unemployment switching mechanism, and exponentially decreasing ($ExpD$) weights $w_1 > \cdots > w_d$ are used. This is consistent with partial-adjustment models of unemployment or delays in firms' responses to demand shocks. The second

[5] Rothman (1998) employs AR, (S)TAR, and bilinear models for predicting U.S. unemployment; Parker and Rothman (1998) apply Beaudry and Koop's (1993) current-depth-of-recession approach; Skalin and Teräsvirta (1998) use STAR models; and Franses and Paap (1998) introduce and apply AR models with censored latent-effect parameters.

[6] The average transition variable of Koop and Potter (1999) also yields qualitatively similar regimes to those produced by the *LD* in an M-TAR model.

gives more weight to distant unemployment changes using exponentially increasing (*ExpI*) weights $w_1 < \cdots < w_d$. It can capture the supply side (discouraged-worker effect) and demand side (possible employer discrimination against the long-term unemployed) impact of unemployment. The resulting variables are denoted $q_t(ExpD, d)$ and $q_t(ExpI, d)$, respectively.[7] To avoid having to take a position on the most adequate switching variable, which could vary by country and sample, we let the data decide using the least-squares principle. This is discussed below.

3.2.1 Identification and estimation issues The Granger (1993) and Teräsvirta (1994) "specific-to-general" strategy for building nonlinear time series models is utilized. At the outset, the linear AR model nested in (3.3) for $\rho_1 = \rho_2 = \rho$ is specified, and an adequate k is selected by a testing down procedure from $k_{\max} = 18$ and corroborated by the Akaike information criterion (AIC). The next stage is to generalize this AR specification to an M-TAR whose parameter vector $(\mu, \rho, \beta, \sigma^2)'$, where $\rho = (\rho_1, \rho_2)'$ and $\beta = (\beta_1, \ldots, \beta_k)'$, is estimated as follows.

Assume the threshold variable $q_t(\mathbf{w}, d)$ is determined a priori.[8] To allow for a time-varying attractor for u_t, the sequence $\{u_t\}_{t=1}^T$ is regressed on the deterministic components $a_0 + a_1 t$ (or on a_0 only), and model (3.3) is fitted to the residual series.[9] Rather than taking the sample mean of this demeaned (and detrended) unemployment series as an estimate for its attractor μ, this is estimated by a grid search. The motivation is that for asymmetric processes the sample mean is a biased estimate of the level of the attractor. The remaining parameters are estimated sequentially by ordinary least squares (OLS) conditional on each grid point yielding $\hat{\rho}(\mu)$, $\hat{\beta}(\mu)$, and residual variance $\hat{\sigma}^2(\mu)$ for $\mu \in \Gamma$, where Γ is a set of plausible values.[10] The attractor estimate is obtained by

$$\hat{\mu} = \arg\min_{\mu \in \Gamma} \hat{\sigma}^2(\mu)$$

and the estimates of the remaining parameters are given by $\hat{\rho}(\hat{\mu})$ and $\hat{\beta}(\hat{\mu})$.

Effectively, this modeling approach implies that the steady state for the unemployment sequence $\{u_t\}_{t=1}^T$ is proxied by $\hat{\mu} + (\hat{a}_0 + \hat{a}_1 t)$. Under the assumption of Gaussian white-noise innovations, this sequential (conditional) OLS procedure is statistically equivalent to maximizing the likelihood function. Moreover, the asymptotic theory for continuous TARs developed by Chan and Tsay (1998) applies to (3.3). It follows that, under the assumptions of stationarity and geometric ergodicity, and i.i.d. but not necessarily Gaussian innovations, conditional OLS yields estimators $(\hat{\mu}, \hat{\rho}, \hat{\beta})$ that are almost surely consistent at rate $\sqrt{T}$ and asymptotically normal.

3.2.2 Threshold unit root and linearity tests The M-TAR framework facilitates tests for the hypothesis of unemployment hysteresis in the context of business cycle asymmetries. More particularly, this paper deploys the Coakley and Fuertes (2000) bootstrap likelihood ratio (LR) test of the restriction $\rho_1 = \rho_2 = 0$ in (3.3) against stationarity with possible M-TAR asymmetry. The latter extends the threshold unit root test introduced by Enders and Granger (1998) by generalizing their threshold variable Δu_{t-1}. If the unit root hypothesis is

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[7] The *ExpI* weights are $w_j = \frac{e^{(j-1)\tau}}{\kappa_I}$, $j = 1, \ldots, d$, where $\tau = \frac{1}{d-1}$ and $\kappa_I = \sum_j e^{(j-1)\tau}$ is a normalizing factor such that $\sum_j w_j = 1$. For *ExpD*, $w_j = \frac{e^{1-(j-1)\tau}}{\kappa_D}$, $\kappa_D = \sum_j e^{1-(j-1)\tau}$.

[8] If the threshold variable is unknown, as is common in practice, the following M-TAR estimation procedure can be repeated for $q_t(LD, d)$, $q_t(ExpI, d)$ and $q_t(ExpD, d)$, with $d = \{1, \ldots, D\}$. Using the least-squares principle, the best-fit threshold variable $\hat{q}_t$ is determined by minimizing the residual variance of the estimated models.

[9] To decide on the inclusion or otherwise of $a_1 t$, we tested for the significance of a deterministic linear trend in the AR filter. See Kapetanios (1999) for a discussion of trends in TAR models.

[10] Our grid search for μ is conducted between the .10($\tau_{.10}$) and .90($\tau_{.90}$) quantiles of the demeaned (and detrended) series with step size $\lambda_G \leq \frac{\tau_{.90} - \tau_{.10}}{m}$, where m is a positive integer to allow for at least G grid points. We use $G = 100$. This approach yields reasonably accurate estimates, as shown in Coakley, Fuertes, and Pérez (forthcoming).

falsified, the next stage is to test for the symmetric adjustment restriction $\rho_1 = \rho_2 = \rho$ against M-TAR asymmetries by means of another bootstrapped LR statistic.[11]

3.2.3 Nonlinear impulse responses Impulse response functions are estimated to shed light on the dynamic behavior of the M-TAR models.[12] We rely on the nonlinear theory developed by Koop, Pesaran, and Potter (1996) and Potter (2000), which extends the traditional analysis. Although for linear time series, several definitions of impulse response functions in the literature are informationally equivalent, they all contain different information in the nonlinear case. Koop, Pesaran, and Potter (1996) and Potter (2000) resolve this problem indirectly by highlighting the superiority of the generalized impulse response (GIR) function. By conceptualizing the GIR as a random variable on the underlying probability space of the time series, a Monte Carlo simulation technique can be employed to approximate the underlying conditional expectations as follows.

For each combination of history $h_{t-1} = \{u_{t-1}, u_{t-2}, \ldots\}$ and time t shock ε_t, which act as initial conditions, use the estimated M-TAR parameters and N randomly selected (future) shocks $V_t = \{v_{t+1}, \ldots, v_{t+N}\}$ to generate R sets of forecasts for the *shocked* system, $\{u_{t+n}(h_{t-1}, \varepsilon_t)\}_{n=0}^{N}$. Generate R sets of *baseline* forecasts $\{u_{t+n}(h_{t-1})\}_{n=0}^{N}$ using a randomly sampled time t shock, v_t, and the same history and random future shocks as in the earlier forecasts. The history- and shock-specific GIR_u is defined as

$$GIR_u(n, h_{t-1}, \varepsilon_t) = E[u_{t+n} \mid \varepsilon_t, h_{t-1}] - E[u_{t+n} \mid h_{t+1}], \;\; n = 0, 1, \ldots, N$$

and can be approximated by averaging the difference between the two types of forecasts over the R replications. This is repeated M times for different combinations of history and shock, and the theoretical GIR function is estimated by averaging the M individual draws.

To measure the speed at which the effect of a shock is absorbed, we employ the half-life (HL) or the half-absorption time. This is defined as the time that must elapse for half of the initial impact to disappear. The history- and shock-specific HL is computed as

$$HL_u = \{m\colon\ |GIR_u(n, h_{t-1}, \varepsilon_t) - GIR_u^{\infty}| \leq 0.5|\varepsilon_t - GIR_u^{\infty}|, \forall n \geq m\}$$

where GIR_u^{∞} denotes the eventual response, or $GIR_u(n, h_{t-1}, \varepsilon_t)$ as $n \to \infty$, which equals zero for mean-reverting processes.

4 Empirical Analysis

4.1 Hysteresis and structural breaks

Total monthly unemployment rates are taken from Datastream for three economies chosen to represent a range of diverse unemployment institutions and experiences. For the United Kingdom and the United States the data span is 1960:1–1999:2, and for Germany the coverage is 1962:2–1999:6. Table 1 summarizes the data.[13] Although all three series have a comparable mean over the sample period, the two European samples show five to six times more dispersion. In all three series the Jarque-Bera (JB) test suggests nonnormality, possibly reflecting structural breaks, nonlinear dynamics, or both.

The variable analyzed is the unemployment rate, even though this is bounded both above and below. This follows Caner and Hansen (1998), who employed a related TAR specification for the U.S. unemployment rate and found that their empirical results were invariant to four transformations of the dependent variable that are

[11]The possibility of a no-adjustment band around the attractor, which is popular in the modeling of financial variables (Coakley and Fuertes, 2001), is implicitly ruled out here.
[12]We are grateful to an anonymous reviewer for suggesting this analysis.
[13]Ideally one would have liked to analyze seasonally unadjusted data, but only adjusted data were available at a monthly frequency for all three countries.

Table 1
Summary statistics for unemployment

Country	Period	T	Mean	Var.	Min.	Max.	S	$K-3$	JB
United States	1960:1–99:2	470	6.05	2.22	3.40	10.80	0.55	0.37	26.48*
United Kingdom	1960:1–99:2	470	5.38	11.01	1.20	11.20	0.41	−1.33	47.79*
Germany	1962:1–99:6	450	5.63	13.68	0.40	13.10	0.14	−1.37	36.60*

Note: T denotes sample size, S and K are skewness and kurtosis coefficients, respectively. JB is the Jarque-Bera normality test statistic, asymptotically distributed as $\chi^2_{(2)}$.
*Significant at the 1% level.

unbounded in one or both directions.[14] The augmented Dickey-Fuller (ADF) procedure with a constant (and trend) applied to the full samples yields the following $\tau_c(\tau_{ct})$ test statistics: −2.70 (−2.64), −2.20 (−2.62), and −0.68 (−2.80), for the United States, United Kingdom, and Germany, respectively. These are in line with the conventional wisdom of apparent unemployment hysteresis.

Figure 1 depicts all three series and highlights the contrast between U.S. and European rates emphasized by Røed (1997), inter alios. Whereas the German and U.K. series appear to display the typical European pattern of an inexorable movement from low to high unemployment between the 1960s and 1980s, this pattern is not so apparent in the U.S. case.

The density functions in Figure 2 indicate that all three unemployment series appear to be bimodal at a minimum. Hence, the ZA test is deployed to investigate the conjecture of hysteresis in unemployment against the alternative of stationarity with one structural break. To allow for a change in both the level and growth rate, the ZA test is initially conducted on model (3.1). The standard errors on the dummy variable coefficients, reported in Table 2, suggest that U.K. unemployment exhibits a significant break in both its level and its slope. By contrast, there is a significant shift in mean for both the U.S. and German series, but the slope dummy is clearly insignificant. Hence, the latter is dropped in the U.S. and German models, and the resultant break date estimates are 1973:10 for the United States, 1980:1 for the United Kingdom, and 1980:9 for Germany. These dates are very close to those found by Bianchi and Zoega (1998) for the German and U.K. series and by Papell, Murray, and Ghiblawi (1999) for all three series.[15] The sample means for the pre-(post-)break subsamples are 4.94 (6.65) for the United States, 2.56 (8.34) for the United Kingdom, and 2.12 (9.15) for Germany.[16] These are taken as prima facie evidence that the extent of the structural break was far more substantial for the two European economies than it was for the United States in both absolute and relative terms. In other words, the impact of shocks on the European economies appears to have been far more severe than on the United States.

The identified breaks seem to be preceded by prior changes in underlying macroeconomic variables such as the level of real oil prices (Oswald, 1999), the ex ante real rate of interest (Phelps, 1994), and the rate of technical progress (Pissarides, 1990; Aghion and Howitt, 1992; Hoon and Phelps, 1997). The increase in oil prices in 1973 and 1979 may be related to all three structural breaks. The break dates for Germany and the United Kingdom appear consistent with the jump in world real interest rates in the early 1980s as a result of disinflationary policies. Moreover, a productivity slowdown occurred around the 1973 break in the United States and some years later in Europe. We should emphasize, however, that this is suggestive only. Establishing a causal link from the changes in the (nonmonetary) fundamentals to the estimated break dates for each of the three economies' natural rates lies beyond the scope of this paper.

[14]Koop and Potter (1999) use a logistic transformation to tackle this problem, which was among the four transformations considered by Caner and Hansen (1998).

[15]The timing of the U.K. break is also virtually the same as that for the single break found by Haldane and Quah (1999) in their analysis of postwar U.K. Phillips curves.

[16]The stubbornly high German rate post-1990 may well reflect the effects of reunification rather than a lack of adjustment and thus may not be typical of other European economies.

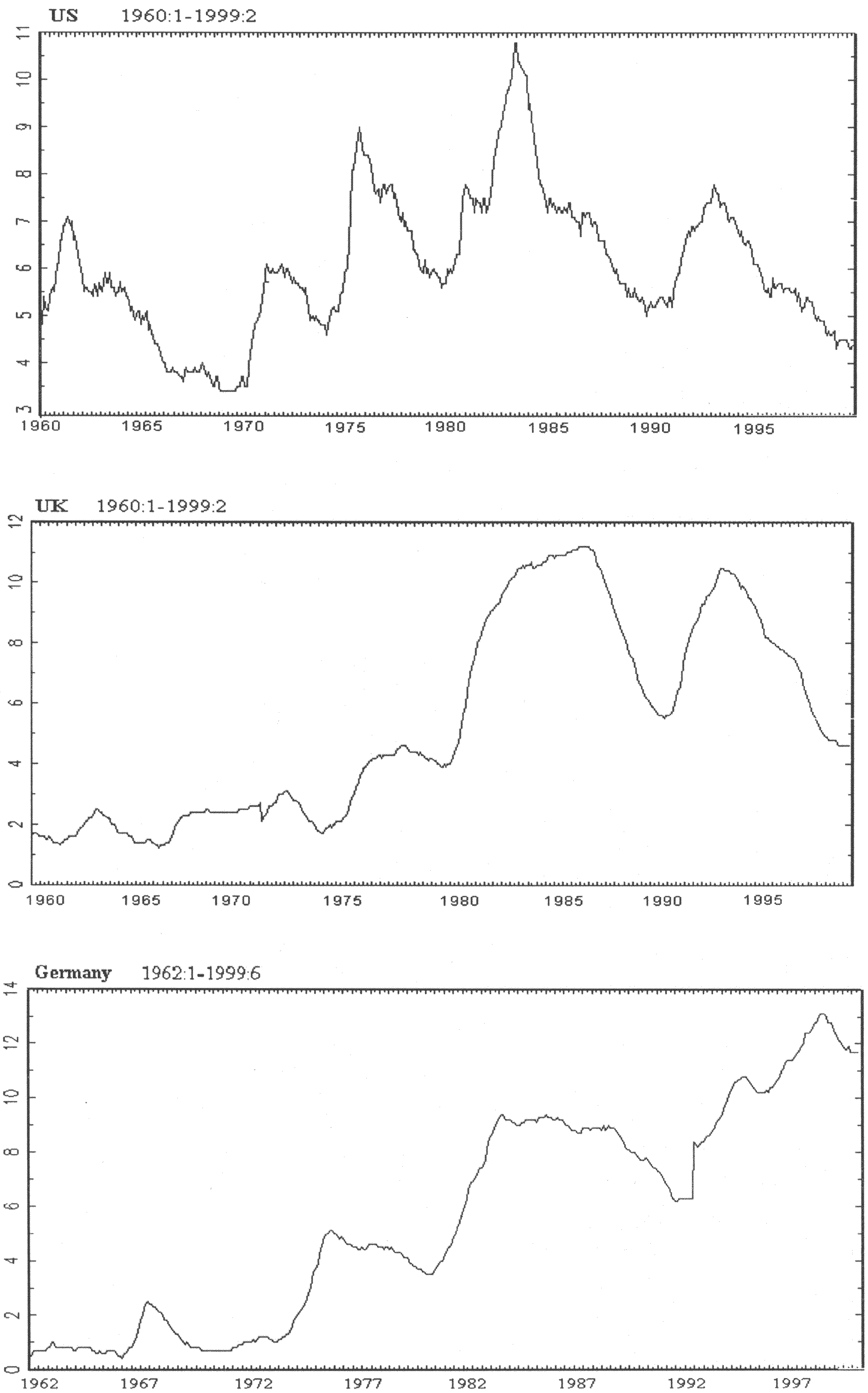


Figure 1
Unemployment rates (%).

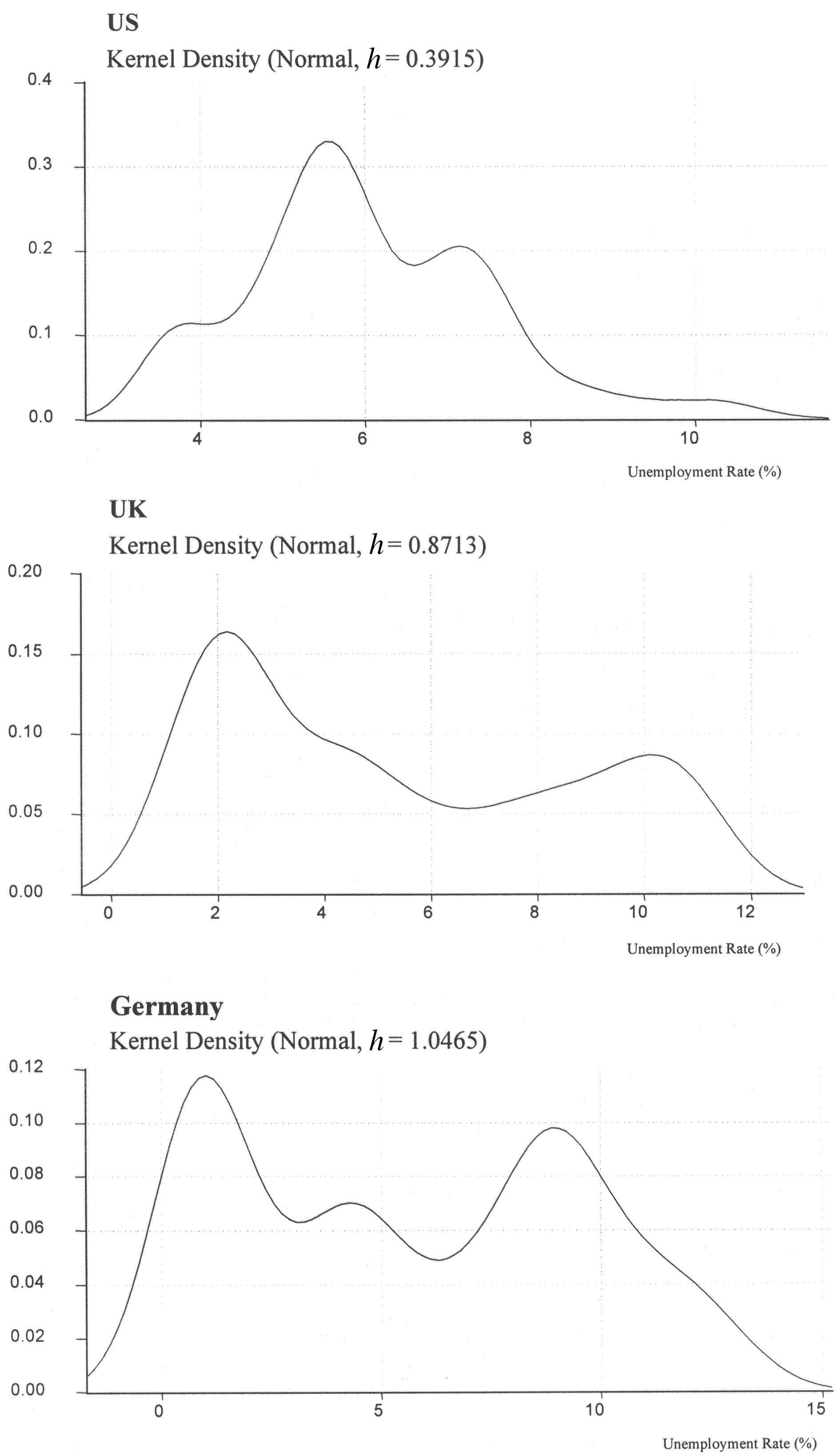


Figure 2
Unemployment density functions.

Table 2
Zivot and Andrews (1992) sequential unit root test

Country	$\hat{\theta}_0$	$\hat{\theta}_1$	$\hat{\theta}_2$	$\hat{\theta}_3$	$\hat{T}_B$	k	$\min_\gamma t(\gamma)$
United States	0.144	3.8×10^{-5}	0.111	-4.1×10^{-4}	1974:5	4	−4.221
	(0.050)	(2.9×10^{-4})	(0.041)	(3.1×10^{-4})			
	0.163	-2.9×10^{-4}	0.127	—	1973:10	4	−4.195
	(0.042)	(1.1×10^{-4})	(0.037)				
United Kingdom	0.015	1.8×10^{-4}	0.092	-4.8×10^{-4}	1980:1	5	−5.415*
	(0.011)	(8.1×10^{-5})	(0.021)	(1.3×10^{-4})			
Germany	−0.021	5.9×10^{-4}	−0.073	9.2×10^{-4}	1992:1	4	−3.488
	(0.016)	(1.7×10^{-4})	(0.035)	(6.1×10^{-4})			
	0.005	3.6×10^{-4}	0.059	—	1980:9	4	−3.422
	(0.015)	(1.6×10^{-4})	(0.028)				

Note: ADF(k) model with deterministic trend function $\theta_0+\theta_1 t+\theta_2 DU_t+\theta_3 DT_t$ (level and slope dummies). Standard errors in parentheses. $\gamma = T_B/T$ with $T_B \in [20, T-20]$.
*Significant at the 5% level using the asymptotic critical values in Zivot and Andrews (1992).

As Table 2 shows, the earlier ADF findings are overturned by the ZA statistic for the U.K. series at the 5% level. This implies that a sharp rise in the equilibrium rate in the early 1980s can explain a good deal of the high U.K. unemployment during the last two decades and provides support for the structuralist school. However, the ZA statistic is clearly nonsignificant at even the 10% level for the other two series. This may stem from the restrictive alternative hypothesis, which permits just one break. Although in principle one could allow for two or more breaks until the null is rejected, this paper seeks to pursue an alternative line of enquiry. This is motivated by our initial conjecture that a single regime shift in conjunction with business cycle asymmetries might be able to explain the apparent $I(1)$-ness of unemployment.[17]

4.2 Asymmetric adjustment

A range of portmanteau nonlinearity tests is applied to detect potential asymmetric behavior in the unemployment series. The tests include the Randles et al. (1980) triples test, the Brock, Dechert, and Scheinkman [BDS] (1987) test, the Keenan (1985) test, the Ramsey RESET1 (1969) and Thursby and Schmidt RESET2 (1977) tests, and the Luukkonen, Saikkonen, and Teräsvirta [LST] (1988) augmented F-test.[18] With the exception of the former, which is applied to the first-differenced series, all tests are computed on the residuals of an AR filter with break dummies at the dates identified in Section 4.1.

The BDS test is designed to reveal hidden patterns that should not occur in truly i.i.d. data. It builds on the correlation integral that measures the fraction of pairs $(\mathbf{y}_t^m, \mathbf{y}_s^m)$ that are within a distance ϵ from each other, where the m-dimensional vector $\mathbf{y}_t^m = (y_t, y_{t+1}, \ldots, y_{t+m-1})$ represents an "m -history" of the data after removing any linear dependence. Under the i.i.d. null, the BDS statistic is asymptotically distributed as a standard normal. As Brock, Hsieh, and LeBaron (1991) and Brooks (1999) inter alios show, however, sample size is an important issue in obtaining reliable inference results. To circumvent this problem, the empirical distribution function of the BDS(ϵ, m) statistic is computed under the null following the bootstrap procedure described in Berkowitz and Kilian (2000, Section 2.1).[19]

[17] Methods related to those developed in Lumsdaine and Papell (1997) could, in principle, be used to test for $I(1)$-ness in the presence of multiple breaks. A single break, however, is consistent with the evidence in Bianchi and Zoega (1998) from a multimodality test that suggests two modes for a number of European economies. In addition, Papell, Murray, and Ghiblawi (1999) show for 16 OECD unemployment series that, once a single break is introduced, there is a dramatic reduction in persistence, whether or not the unit root is rejected.

[18] For an overview of these tests, see Brock, Hsieh, and LeBaron (1991) and Granger and Teräsvirta (1993). The triples test was introduced to the economics literature by Verbrugge (1997).

[19] We are grateful to an anonymous reviewer for this suggestion. The empirical distribution function is quite close to the $N(0, 1)$ cumulative distribution function for all (ϵ, m) cases considered. This is consistent with Brooks's (1999) results for $T = 500$, which is close to our sample sizes.

Table 3
Portmanteau nonlinearity tests

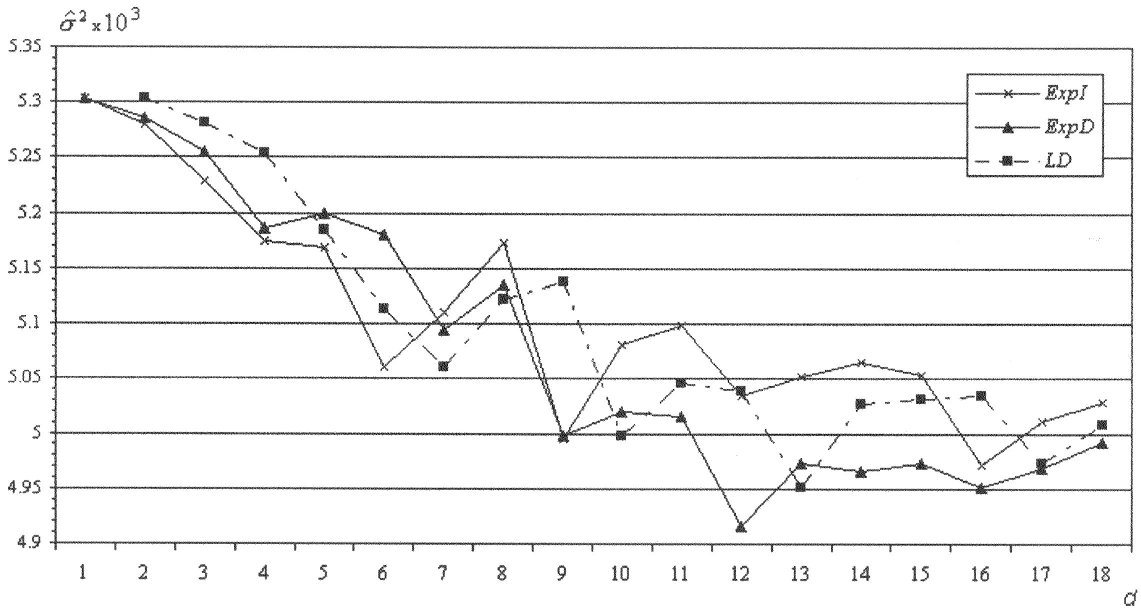
	I. BDS test				
	ϵ/σ	m			
Country		2	3	4	5
United States	.75	1.285	2.039	2.304	2.343
		[0.090]	[0.030]	[0.009]	[0.031]
	1	1.620	2.135	2.196	1.981
		[0.065]	[0.017]	[0.028]	[0.035]
	1.25	2.160	2.403	2.565	2.440
		[0.019]	[0.009]	[0.012]	[0.005]
United Kingdom	.75	2.105	2.378	2.644	2.251
		[0.029]	[0.018]	[0.011]	[0.030]
	1	4.369	4.817	4.822	4.302
		[0.000]	[0.000]	[0.000]	[0.000]
	1.25	2.973	3.581	3.526	3.096
		[0.004]	[0.002]	[0.000]	[0.001]
Germany	.75	4.435	4.668	4.426	4.096
		[0.002]	[0.000]	[0.003]	[0.004]
	1	5.043	5.187	4.823	4.464
		[0.002]	[0.000]	[0.001]	[0.002]
	1.25	4.999	4.893	4.412	3.891
		[0.001]	[0.002]	[0.003]	[0.003]
	II. Other tests				
	Triples	Keenan	RESET1	RESET2	LST
United States	1.869	4.550	4.362	6.276	1.314
		[0.033]	[0.002]	[0.000]	[0.189]
United Kingdom	1.796	3.539	1.133	9.562	2.076
		[0.061]	[0.340]	[0.000]	[0.006]
Germany	4.963*	6.029	3.992	8.325	1.997
		[0.014]	[0.003]	[0.000]	[0.014]

Note: Bootstrap p-values in brackets for the BDS statistic. The triples test is based on asymptotic $N(0, 1)$ critical values. For the remaining tests asymptotic p-values are reported in brackets.
*Significant at the 5% level.

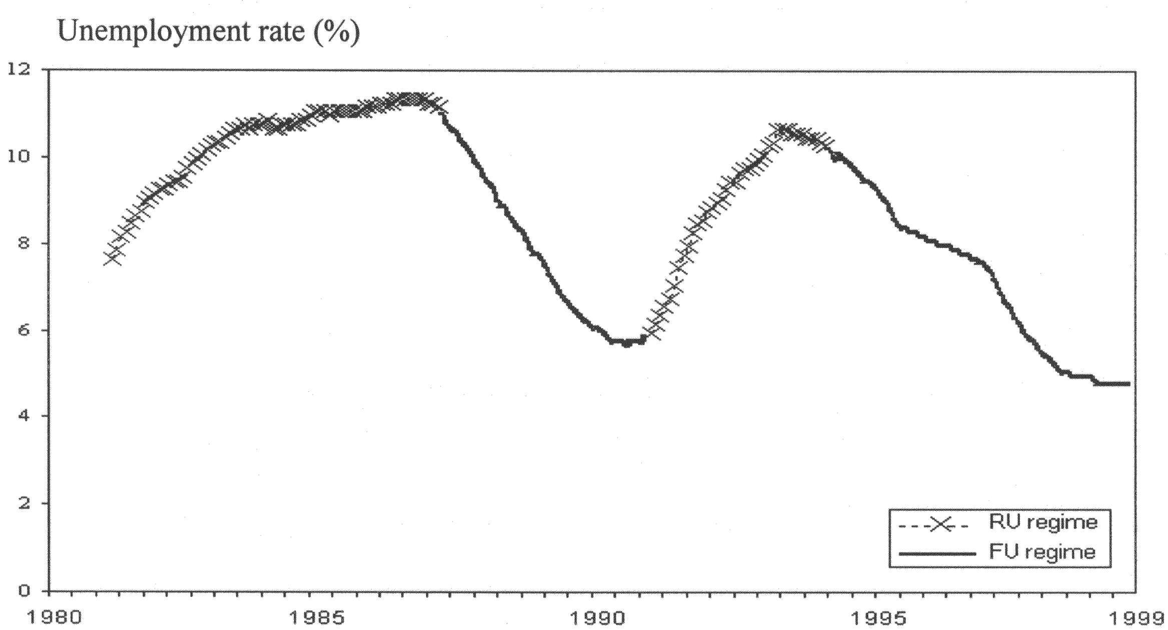
Table 3 reports the results. The BDS bootstrap p-values from 999 replications suggest nonlinear unemployment dynamics for the three series. This confirms the BDS test findings in Açemoglu and Scott (1994) for U.K. unemployment, albeit for a shorter sample span. The triples test supports the conjecture of steepness (or asymmetries in growth rates) for Germany at the 1% level and for both the United Kingdom and United States at the 10% level. The other four test statistics, with the exception of the LST statistic for the United States and the RESET1 statistic for the United Kingdom, are all significant, at better than the 5% level in most cases.

The M-TAR model (3.3) is estimated for the subsamples obtained by splitting each unemployment series at the identified break. Figure 3a illustrates the exercise of finding the best-fit threshold variable for the United Kingdom 1980–1999. It plots the residual variance of the M-TAR models for the three threshold variables considered using $d = \{1, \ldots, 18\}$. The best-fit switching variable is $q_t(ExpD, 12)$. A similar procedure is used for the other subsamples and countries. The threshold variables $q_t(ExpI, d)$ and $q_t(ExpD, d)$ yield the best fit in five out of the six subsamples. Table 4 reports the M-TAR parameter estimates.[20] The identified d spans at least 1 year (with one exception), and thus the state variable $q_t(\mathbf{w}, d)$ may be taken to represent a sustained period of past rising or falling unemployment. The RU regime embraces the range from just beyond an

[20]Our proxy for u_t^*, obtained as discussed in Section 3.2.1, is $5.72I(t \le 1973{:}10) + (9.18 - .007t)I(t > 1973{:}10)$ for the United States, $(.65 + .012t)I(t \le 1980{:}1) + (11.99 - 0.12t)I(t > 1980{:}1)$ for the United Kingdom, and $(.81 + .020)I(t \le 1980{:}10) + 11.86I(t > 1980{:}10)$ for Germany, where $I(\cdot)$ is an indicator function. These confirm our earlier verdict (from the subsample means) that the regime shift seems less severe in the United States.



(a)



(b)

Figure 3
Estimated residual variance and observations classified by regime (U.K. 1980:2–1999:2).

unemployment low as the economy comes out of a boom to beyond a recessionary peak. Similarly, the FU regime stretches from beyond an unemployment high as the economy comes out of a recession to beyond its trough in a boom. Figure 3b illustrates this by plotting postbreak U.K. unemployment classified by regime.

The estimated adjustment speed for the RU regime, ρ_1, is absolutely larger—implying more rapid dynamics—than that in the FU regime, ρ_2, which is insignificant at the 5% level in all cases.[21] This provides prima facie evidence of business cycle asymmetries. The fast-up dynamics in the RU regime are most pertinent when unemployment is emerging from the low of an economic boom and rising toward its equilibrium level as the economy slows down. By contrast, the pull from the attractor is weak throughout the FU regime, which is characterized by slow-down dynamics. One clear implication is that unemployment tends to rise more quickly in recessions than it declines in expansions. In other words, unemployment is more sluggish in the aftermath of a recession.

A test for unit root behavior against M-TAR dynamics is then implemented. Bootstrap p-values are derived by resampling using the null model estimates and normally distributed random disturbances with variance equal to the estimated residual variance. The p-values reported in Table 5 falsify the null at the 2% level or better for both subsamples in all three economies. The implication is overall mean reversion despite local persistence. These findings contrast with the ADF test results for subsamples.[22]

Another LR statistic is used to test for symmetric adjustment, $\rho_1 = \rho_2$, against distinct dynamics in the FU and RU regimes. For an asymptotically pivotal statistic such as this, the bootstrap test size distortion is at least $T^{-1/2}$ smaller than that of the corresponding asymptotic χ^2 test. This motivates computing bootstrap p-values, which, as evidenced in Table 5, reject symmetry at better than the 5% in most cases.[23] Hence, business cycle asymmetries are germane to a proper understanding of unemployment persistence.

A final issue is whether cyclical asymmetries on their own can explain the apparent unemployment hysteresis. M-TAR models are fitted to the full samples ignoring the structural breaks. By contrast with the results for subsamples, the unreported threshold unit root test rejects the null only for the United States at the 5% level. This, together with our findings in Section 4.1, suggests that ignoring either cyclical asymmetries or structural breaks (or indeed both) gives an exaggerated picture of unemployment persistence.

4.2.1 Propagation of shocks Since it is difficult to evaluate the dynamic properties of an M-TAR model from its parameter estimates, GIRs are computed to assess how shocks propagate. The analysis is conditioned on the RU and FU regimes. Thus, one GIR is computed for each regime by randomly selecting histories (with replacement) from the observed time series, such that $q_t(\mathbf{w}, d) \geq 0$ and $q_t(\mathbf{w}, d) < 0$, respectively. The future shocks (and time t shocks for the baseline forecasts) are randomly drawn with replacement from the estimated residual vector. The current shock is defined as $\varepsilon_t = \delta\hat{\sigma}$, where $\hat{\sigma}$ is the M-TAR regression standard error and $\delta = \pm 1$ and ± 2 to explore the effect of "small" and "large" shocks, respectively. We draw 100 histories for each regime and conduct 500 Monte Carlo replications for each history. The maximum horizon of the GIR is set to 120 months, and all GIRs are normalized so that the initial effect of the shock is δ for all histories.

Figure 4 plots the regime- and shock-dependent GIRs for each subsample, and Table 6 reports the associated HL estimates. What are the main findings? First, the overall HL of all unemployment shocks across

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[21] The unreported Eitrheim and Teräsvirta (1996) Lagrange multiplier diagnostic tests—deployed for (3.3) using a transition parameter equal to 10^3—indicate that the M-TAR specification is plausible, as there is no evidence of autocorrelation or remaining nonlinearity at the 5% level. The latter is corroborated by the BDS test on the M-TAR residuals, suggesting that dynamic asymmetries account for most of the nonlinear dependence in unemployment.

[22] These yield only two significant statistics (out of six) at the 5% level: United States post-1973:10 and United Kingdom post-1980:1. The overall (both subsamples) mean reversion conclusion from the M-TAR test also contrasts with the evidence from the ADF and ZA tests for full samples.

[23] As a check for robustness, the Hodrick-Prescott (HP) detrended series are split at the identified structural break and the M-TAR asymmetry tests are applied. Despite the smoothing effects of the HP filter resulting in a decrease in asymmetry as shown in Psaradakis and Sola (1997), the p-values still indicate significant asymmetries. Note that this is subject to the caveat that the break points may change following HP detrending. We are grateful to an anonymous reviewer for pointing this out.

Table 4
M-TAR models

				$q_t(\mathbf{w}, d)$		RU regime		FU regime						
Country	Sample	k	μ	$\mathbf{w}$	d	ρ_1	$\%T$	ρ_2	$\%T$	Higher-order terms $\beta_1, \beta_2, \ldots, \beta_k$				
United States	1960:1–73:10	4	0.78	ExpI	15	−0.079	34	−0.015	66	−0.16	0.13	0.16	0.20	—
						(0.022)		(0.014)		(0.080)	(0.079)	(0.077)	(0.078)	
	1973:11–99:2	4	1.43	ExpI	16	−0.046	39	−0.011	61	−0.03	0.13	0.17	0.17	—
						(0.013)		(0.010)		(0.058)	(0.056)	(0.054)	(0.055)	
United Kingdom	1960:1–80:1	5	−0.52	LD	13	−0.033	43	−0.018	57	0.09	0.25	0.14	0.07	0.16
						(0.012)		(0.013)		(0.065)	(0.066)	(0.067)	(0.067)	(0.066)
	1980:2–99:2	3	2.27	ExpD	12	−0.024	50	−0.004	50	0.12	0.35	0.27	—	—
						(0.006)		(0.003)		(0.066)	(0.061)	(0.064)		
Germany	1962:1–80:9	3	0.93	ExpD	14	−0.019	30	−0.011	70	0.23	0.25	0.30	—	—
						(0.008)		(0.006)		(0.065)	(0.065)	(0.064)		
	1980:10–99:6	3	2.71	ExpI	8	−0.033	56	0.004	44	0.022	0.025	0.091	—	—
						(0.008)		(0.006)		(0.067)	(0.068)	(0.054)		

Note: M-TAR models fitted to the demeaned series, or to the demeaned and detrended series if the deterministic time trend in an AR filter is significant. An intercept was included in the M-TAR models because the attractor may differ from the mean and hence the residuals may have a nonzero mean. Standard errors in parentheses. $\%T$ denotes the percentage of observations in each regime.

Table 5
Bootstrap LR threshold unit root and symmetry tests

Country	Sample	Unit root test	Symmetry test	
			u_t	HP-detrended u_t
United States	1960:1–73:10	34.84 [0.000]	32.99 [0.000]	24.27 [0.018]
	1973:11–99:2	52.97 [0.000]	41.70 [0.000]	43.38 [0.000]
United Kingdom	1960:1–80:1	24.85 [0.020]	16.74 [0.058]	13.75 [0.115]
	1980:2–99:2	37.56 [0.000]	26.20 [0.001]	29.29 [0.008]
Germany	1962:1–80:9	31.76 [0.012]	20.27 [0.019]	27.00 [0.013]
	1980:10–99:6	24.47 [0.018]	15.83 [0.023]	8.19 [0.218]

Note: Empirical p-values from a parametric bootstrap simulation with 999 replications in brackets.

Table 6
Half-lives of shocks in M-TAR models

Country	Sample	RU regime				FU regime			
		+2	+1	−1	−2	+2	+1	−1	−2
United States	1960:1–73:10	19.41	23.24	24.11	24.51	34.83	35.57	34.75	34.76
	1973:11–99:2	32.57	33.87	33.63	34.71	41.54	41.98	41.95	41.19
United Kingdom	1960:1–80:1	28.44	28.26	27.86	28.63	28.70	29.17	29.60	29.88
	1980:2–99:2	46.06	92.86	88.69	93.77	108.12	107.43	108.11	104.12
Germany	1962:1–80:9	36.01	36.37	36.58	37.10	61.07	62.88	63.36	63.40
	1980:10–99:6	33.50	55.55	65.00	75.06	65.05	—	—	106.5

Note: The shocks are defined as $\varepsilon_t = \delta \cdot \hat{\sigma}$ where $\hat{\sigma}$ is the standard error of the M-TAR model and $\delta = \pm 2$ (large shocks) and $\delta = \pm 1$ (small shocks).

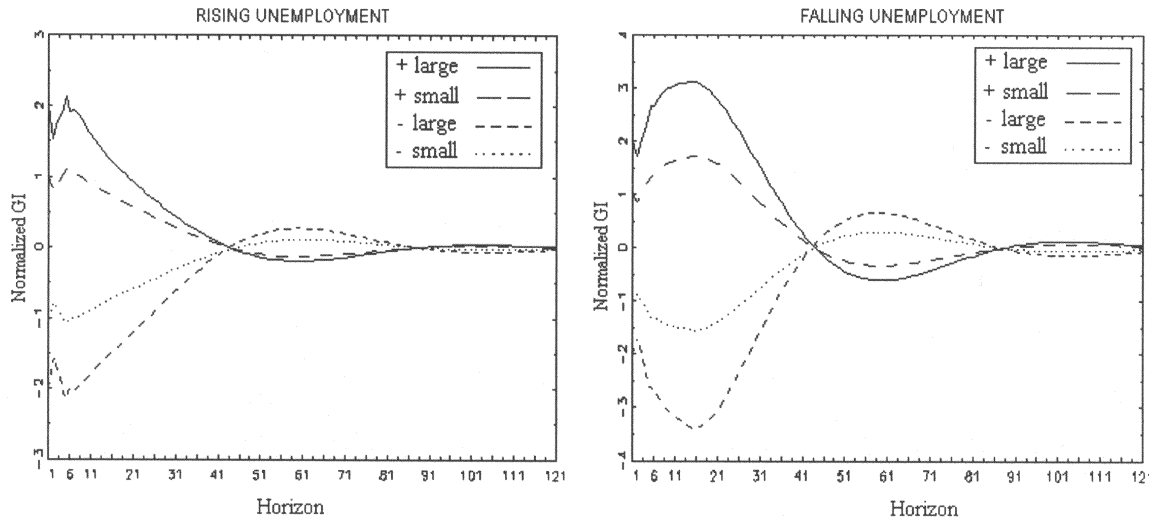
subsamples and regimes reveals an interesting pattern. Shocks are most persistent in Germany, with an average HL of around 5.5 years, followed by just over 5 years for the United Kingdom and over 2.5 years for the United States. This confirms the stylized impression of Germany as the most sclerotic economy and of the United States as the most flexible. The overall picture for the United Kingdom, which seems closer to the German pattern, disguises the fact that the HLs for both the United States and United Kingdom are virtually identical at 29 months in the first subsample, whereas the corresponding HLs for Germany are more persistent. Thus the sharp contrast between U.S. and U.K. labor markets seems entirely a product of recent decades and of how their respective economies responded to adverse shocks, such as the oil price hikes in the 1970s.

Second, although the time profile of U.S. shocks suggests more tardiness in recent as compared with earlier decades, this increase in persistence fades by comparison with that of European labor markets. Shocks to European unemployment in the recent subsample persist much longer and are between two and two-and-a-half times more persistent than their U.S. counterparts. For the United Kingdom and Germany, post-1980 shocks are notably more persistent than in the pre-1980 period. For instance the GIRs reveal that the average HL of shocks to U.K. unemployment is some 2.84 and 3.65 times that of pre-1980 for the RU and FU regimes, respectively. This increase in persistence helps to explain the recent high European unemployment rates and provides support for the persistence school.[24]

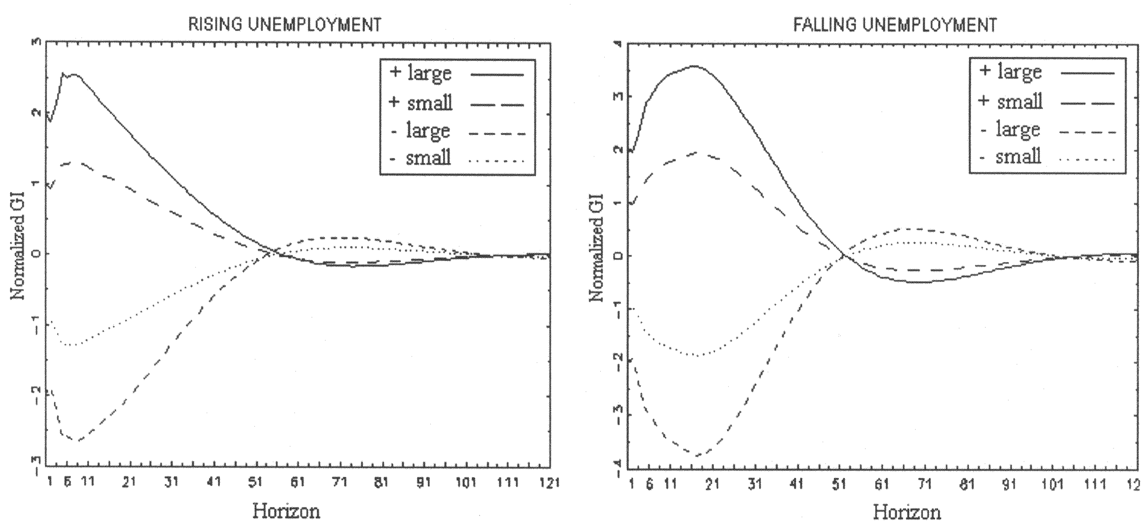
Finally, the FU regime displays high persistence and an exacerbated oscillatory pattern compared to the RU regime.[25] This is especially pronounced for small shocks hitting the FU regime for the (postbreak) German series whose HLs exceed 10 years. Large shocks conditional on the FU regime display marked overshooting behavior in the United Kingdom. The initial impact is magnified by a factor of around five after 20 months, following which it begins to decline. These findings add to the earlier evidence of asymmetries over the business cycle.

[24]A possible rationale for the latter may be found in the chain reaction theory of Karanassou and Snower (1998).

[25]This pattern of more persistent falling unemployment is consistent with that revealed by the GIRs in Altisimo and Violante (2001) from a VAR for output and unemployment that includes a depth-of-recession threshold feedback variable.



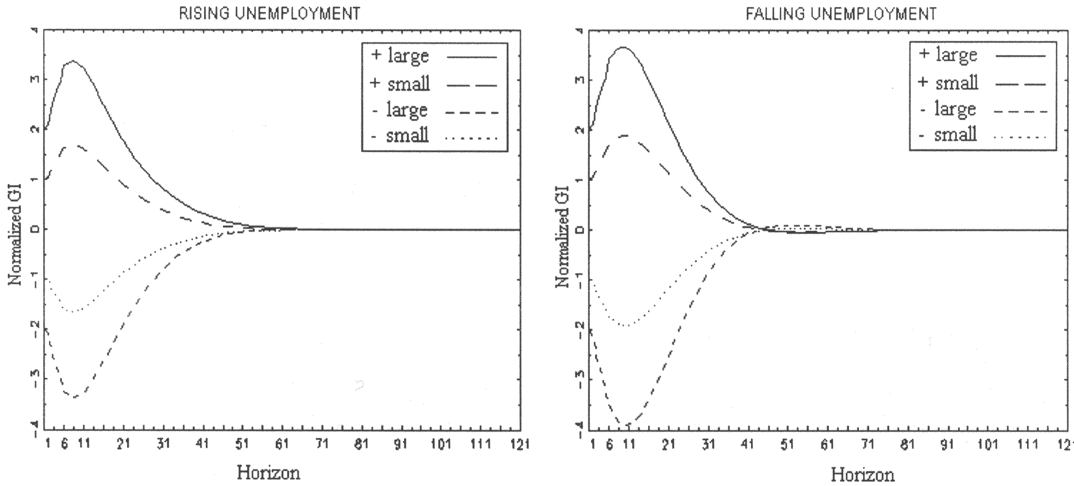
(a) U.S. 1960:1−1973:10



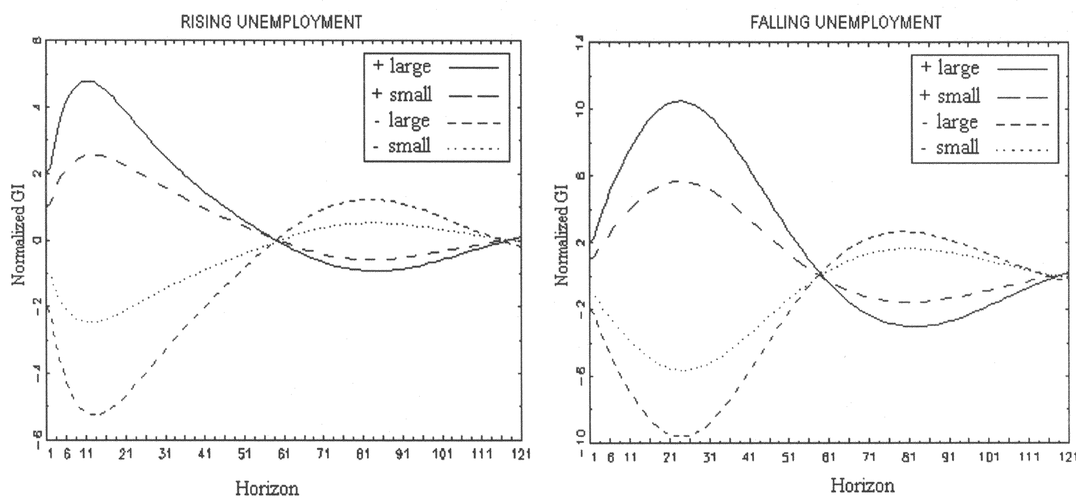
(b) U.S. 1973:11−1999:2

Figure 4
Generalized impulse response functions for M-TAR model.

Figure 4 continued

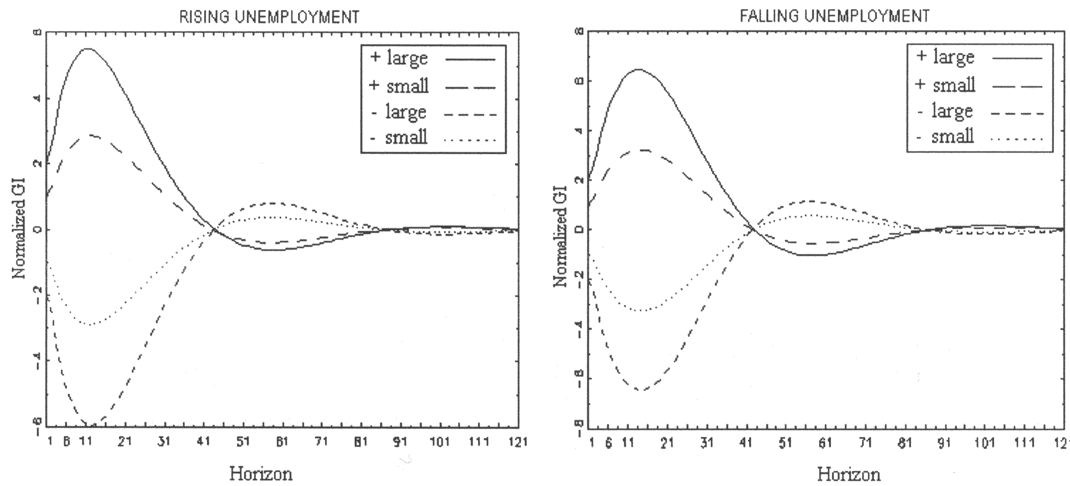


(c) U.K. 1960:1–1980:1

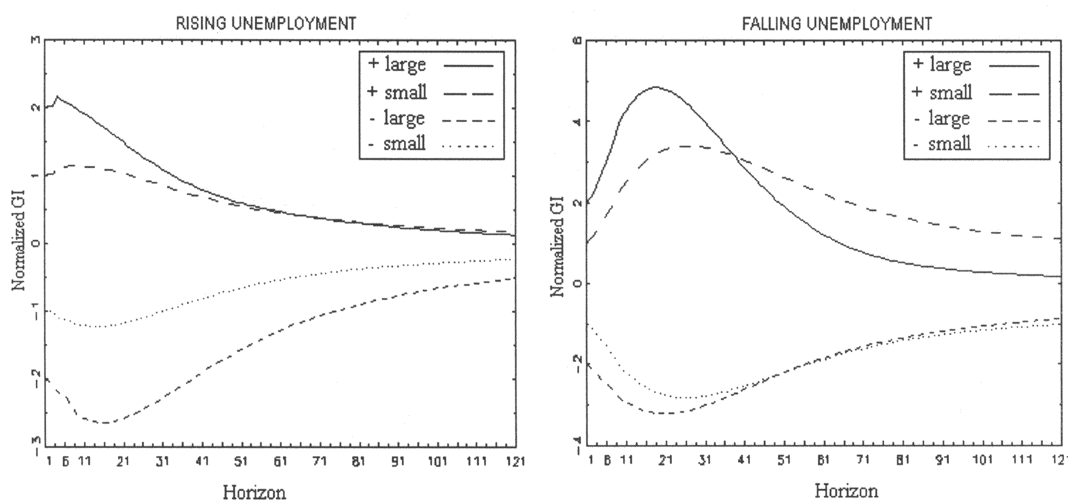


(d) U.K. 1980:2–1999:2

Figure 4 continued



(e) Germany 1962:1–1980:9



(f) Germany 1980:10–1999:6

4.3 Summary and interpretation

The above findings can be summarized as follows. First, average unemployment shifted to a higher plateau in the United Kingdom and Germany in 1980 and in the United States around 1973. The identified break for the United States in 1973, however, was comparatively smaller than that affecting European unemployment, as indicated by the relative change in mean level for the pre- and postbreak subsamples. Since many of the presumed fundamentals of the natural rate—technical progress, real interest rates, oil prices—also shifted around these dates, these findings are in line with structuralist school principles and are also consistent with the results of Bianchi and Zoega (1998) and Papell, Murray, and Ghiblawi (1999).

Second, M-TAR models fitted over subsamples corroborate the fast-up, slow-down dynamics of unemployment over the business cycle. Moreover the GIRs reveal that the HL of shocks has increased markedly over the cycle in the last two decades—from which the persistence school can take some comfort—and that this effect is clearly more pronounced for the United Kingdom and Germany in falling unemployment regimes. This intensified sclerosis of labor markets offers a rationale for the high European unemployment of recent decades, which complements the abrupt change in the natural rate in the early 1980s.

Finally, ignoring either the identified break or dynamic asymmetries produces an exaggerated picture of unemployment persistence. M-TAR models fitted to the full samples (ignoring the break) suggest unit root behavior in U.K. and German employment. The ZA test statistic based on a linear model that allows for a single break is nonsignificant in the U.S. and German series.

5 Conclusions

This paper examines the time-series properties of three unemployment series to assess the persistence and structuralist theories from a nonlinear perspective. The methodology adopted combines a one-time shift in equilibrium unemployment and a nonlinear TAR model capturing business cycle asymmetries. The identified structural breaks suggest an autonomous shift in the natural rate for the United Kingdom and Germany in 1980 and for the United States in 1973, possibly prompted by changes in fundamentals around these dates. The magnitude of the break in relative and absolute terms is substantially larger in the European economies, and this provides support for the structuralist school in explaining the high European unemployment of recent decades.

Nonlinear impulse response functions indicate that shocks are more persistent in the falling- than in the rising-unemployment regime in all economies. More significantly, the half lives of shocks to unemployment have increased sharply postbreak, particularly for the United Kingdom and Germany, which places the persistence school closer to the mark. The conclusion is that elements from both theories are germane to a satisfactory account of medium- to long-term unemployment dynamics in recent decades. Business cycle asymmetries in conjunction with a single structural break seem able to explain the apparent unemployment hysteresis in the unit root sense for all three economies and particularly for Europe.

Some directions for further research can be suggested. Generalizing the M-TAR models to allow for a time-varying attractor parameter subject to occasional breaks at unknown points might be worthwhile. Another possibility would be to build a more general nonlinear model nesting elements from both theories—occasional breaks in the attractor and time-varying speeds of adjustment to equilibrium—to develop tests for their separate and joint significance in explaining the sustained high unemployment of many OECD countries.

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