UNEMPLOYMENT, CYCLE AND GENDER

AMADO PEIRÓ
JORGE BELAIRE-FRANCH
MARIA TERESA GONZALO

FUNDACIÓN DE LAS CAJAS DE AHORROS
DOCUMENTO DE TRABAJO
Nº 581/2011

De conformidad con la base quinta de la convocatoria del Programa de Estímulo a la Investigación, este trabajo ha sido sometido a evaluación externa anónima de especialistas cualificados a fin de contrastar su nivel técnico.

ISSN: 1988-8767

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Amado Peiró*
Jorge Belaire-Franch*
Maria Teresa Gonzalo*

Abstract

This study analyzes the relationship between unemployment and the business cycle in the UK and the US. For both economies, a strong and definite association is found that shows that cyclical shocks extend their effect on unemployment over several quarters. This association is much more intense for male unemployment than for female unemployment, and has lost some strength in the UK in the last few years. Markov switching regime models with two regimes display clear differences between expansions and contractions in both countries.

Key words: Business cycle, gender, Markov switching regime model, unemployment
JEL classification: E32, J64

Corresponding author: Amado Peiró Giménez. E-mail: Amado.Peiro@uv.es

***** Departament of Economic Analysis, University of Valencia

Acknowledgement: The authors are grateful to an anonymous referee and participants at the XXV AIEL Conference of Labour Economics, Chieti-Pescara (Italy), and the 6th Eurostat Colloquium on Modern Tools for Business Cycle Analysis, Luxembourg, for their useful comments and suggestions. Financial support from the Ministerio de Educación y Ciencia (SEJ2007-67101/ECON) is gratefully acknowledged.

1. Introduction

The cyclical nature of economic dynamics is commonly accepted. Periods of high economic activity, or expansions, are followed by periods of low economic activity, or contractions, in a non-regular sequence. Although the duration of all these periods may be very different, they are recurrent and display certain comovements among important economic variables. Numerous researchers have studied these business cycles for many decades. Burns and Mitchell (1946) was a milestone in the research on this issue with its contribution to the definition and measurement of business cycles, and since then the dating of phases and the extraction of cyclical components of economic variables has attracted the attention of many researchers.

Output and unemployment are two key variables of the business cycle, such that they are taken into account in virtually all the research on cycles and are of maximum importance in macroeconomic performance. Their role in business cycles is clear, but their movements have opposite signs. Output is clearly a procyclical variable, undoubtedly the most defining variable of business cycles phases, and its movements give rise to the different expansions and contractions in many chronologies. On the other hand, unemployment is a clearly countercyclical variable and, consequently, increases in contractions and decreases in expansions. These opposite directions across business cycles should give rise to an inverse relationship between output and unemployment. However, several issues may hide this relationship. Firstly, there may be flows from unemployment to out of the labor force and vice versa that could distort or cause difficulties in studying the link between these two variables. These flows may be induced by phenomena such as 'the added worker effect' or 'the discouraged unemployed.' Secondly, the relationship between output and unemployment may be dynamic instead of fully contemporaneous. In addition to a simultaneous effect, the response of unemployment to cyclical shocks may take some time and, therefore, unemployment could be a lagging indicator of business cycles. In fact, according to NBER business cycle dating for the US, unemployment peaked more than one year after the trough in the last two contractions (15 and 19 months after, respectively.) To complicate things even further, besides this dynamic relationship, the labor market could also anticipate future cyclical movements. Finally, changes in aspects such as employment protection legislation and dismissal costs may affect the relationship or its dynamics (see, for example, Alewell, Schott, and Wiegand, 2009 or Wolfers, 2005.) This possible distribution of the effects of business cycles on unemployment may imply that the effect is relatively weak in each single time period, but spreads over time; unemployment may react to cyclical shocks or policy measures a long time after they took place. Consequently, from a policy perspective, it is important to elucidate the time distribution of unemployment variations triggered by business cycles.

The relationship between unemployment and business cycles may have changed in the last few decades. In fact, many researchers believe that recent business cycles are rather different from the preceding ones. Variability in main economic aggregates has decreased, and contractions have been less frequent and severe (since November 1982 the NBER has dated only two contractions in the US lasting eight months each.) Factors such as the increase in the service sector with respect to industry or agriculture, or a major role of the public sector in modern economies, could be behind these changes, and this weakening of cyclical movements may have altered their linkage with unemployment. Nevertheless, the interest in the relationship between output and unemployment over the business cycle has recently increased in the light of current economic events. The pronounced downturn in economic activity that began in 2008 in many countries has sparked a sharp increase in unemployment. This situation has generated huge interest in this topic, especially in those countries which face high unemployment rates unheard of in many years.

The consequences of cyclical movements on unemployment may also differ by gender. While a great deal of literature exists on gender differences in labor participation and wages, research on gender differences in unemployment and their relationship with the business cycle is much more sparse. Traditionally, it has been widely accepted that the labor supply curve is more elastic among women (Killingsworth, 1983, Blundell and MaCurdy, 1999), and this fact could be the basis of existing differences in unemployment across the business cycle. Nevertheless, several researchers have examined these issues from different perspectives and have come to different conclusions. Clark and Summers (1981) found that the cyclical behavior of employment is not age and gender neutral, as employment of young women was more responsive to cyclical shifts than employment of older women, and the latter was, in his turn, more responsive than employment of similarly aged men. Blank (1989) found a stronger association of changes in employment with changes in GDP for women than for men of the same race. Solon, Barsky and Parker (1994) and Kandil and Woods (2003), however, report empirical evidence that questions the higher elasticity of labor supply among women, though their results do not exclude the possibility of different cyclical behavior in the extensive margins for men and women. Queneau and Sen (2008) consider several theories regarding the dynamics of unemployment over the business cycle, and present evidence of gender differences in unemployment dynamics in three out of eight OECD countries, but the degree of persistence in male and female unemployment rates is relatively low in all the countries under examination.

When considering the relationship of business cycles with unemployment, another important aspect is the possible existence of asymmetries. The presumption that important economic variables present asymmetric behavior over the business cycles has a long tradition in economic thought, which dates back to the pioneering work on business cycles by Mitchell (1927.) In fact, many researchers, including Keynes (1936), have firmly believed that business cycles display marked asymmetries. Much later, in a seminal study, Neftçi (1984) formally tested the asymmetric behavior of unemployment over the business cycle and since then a large number of studies have followed. DeLong and Summers (1986), Hussey (1992), Acemoglu and Scott (1994), Koop and Potter (1999), McKay and Reis (2008), among many others, have also found asymmetries in labor market variables. Though these conclusions are not unanimous, one may conclude with Mittnik and Niu (1994) that 'Although the empirical evidence on business cycle asymmetries is somewhat mixed, there appears to be fairly strong support for asymmetries in unemployment data, while there is somewhat weaker support for aggregate output data.' In fact, McKay and Reis (2008) recently proposed a new business cycle feature: 'contractions in employment are briefer and more violent than expansions but we cannot reject the null of equal brevity and violence for expansions and contractions in output.'

If unemployment displays an asymmetric evolution over time but output does not display clear asymmetries, the source of unemployment asymmetries could lie in the nature of its dependence on output. Several researchers have investigated asymmetries and non-linearities in the relationship between unemployment and cyclical movements from the perspective of Okun's law. Virén (2001) presents evidence of non-linearities in Okun's relationship for 18 out of 20 OECD countries. Huang and Chang (2005) find support of threshold non-linearity for Canada. Using Hamilton's flexible nonlinear inference, Huang and Lin (2006) find clear evidence of nonlinearity between the cyclical components of US unemployment and output. Silvapulle, Moosa and Silvapulle (2004) present evidence of asymmetry in the output-unemployment relationship for the US post-war economy so that the response of unemployment is stronger to negative than to positive cyclical output. Holmes and Silverstone (2006) use the Markov regime-switching model to analyze asymmetries in Okun's law for the US. When testing linearity against nonlinearity in US data, Crespo (2003) concludes the existence of a regime dependent Okun's parameter, which implies that cyclical unemployment is more responsive to changes in negative cyclical output.

The aim of this paper is to examine the association between unemployment and business cycles in two main economies, the United Kingdom and the United States. Special attention will be paid to essential aspects of this relationship, such as differences by gender, changes over time and possible asymmetries. A clear dynamic association is found between cycle and unemployment, but this relationship differs by gender and presents asymmetries. In order to achieve these objectives, Section 2 presents the data used in this study. Section 3 analyzes these relationships from a general perspective, by gender and by period of time. Section 4 examines the empirical evidence looking for possible asymmetries between expansions and contractions. Finally, Section 5 summarizes the main conclusions.

2. Data

Quarterly data on GDP for the UK and the US were collected from *International Financial Statistics*, International Monetary Fund, and from the *Bureau of Economic Analysis*, US Department of Commerce, respectively. They cover the period 1971:1–2008:4 and 1948:1–2008:4, respectively, and are seasonally adjusted. With regard to labor markets, quarterly data on unemployment rates were obtained for the UK from the *UK Office for National Statistics* and for the US from the *Bureau of Labor Statistics*, US Department of Labor. They cover the same periods as GDP and are also seasonally adjusted.

For both countries, GDP presents a clear trend. To obtain the cyclical component, these series must be detrended. The detrending procedure is highly controversial, as different methods give rise to different properties of the resulting cyclical component. Probably, the filter proposed by Hodrick and Prescott (1980 and 1997), HP, is the most widely used. This filter has been applied to the logarithm of GDP with a smoothing parameter equal to 1600, the value proposed by Hodrick and Prescott (1997) for quarterly data. The difference between the original series (GDP in logs) and the trend obtained is the cyclical component, which is shown in Figure 1, panels A and B.

Perhaps the most striking feature of these two graphs is the lower variability or volatility that cyclical GDP displays in both countries in the last few years of the sample (though the last quarter registers a hard fall.) Table 1 shows the standard deviations of cyclical GDP in different periods: the whole sample, the first two thirds of the sample and the last third of the sample. The standard deviation of cyclical GDP is clearly lower in the last few years. It is about one third in the UK and about one half in the US. The same happens with the increments in cyclical GDP (see Figure 1, panels C and D, and Table 1.) This much smoother evolution is a well known fact that has been discussed in detail and could be due to several factors such as better statistics, the growing share of services in GDP or the role of the public sector and has led many researchers to enquire about the drastic change in recent business cycles or even their end.[1] It is also interesting to note in Figure 1, panels A and B, that most extreme values of the cyclical component are positive in the UK and negative in the US. This asymmetry is also shown in Table 2. Although the mean of the cyclical components is zero, the five most extreme cyclical components are positive in the UK, while the nine most extreme are negative in the US. However, this feature disappears almost completely in both countries when examining the series of increments in cyclical components (see Table 2 and panels C and D in Figure 1.)

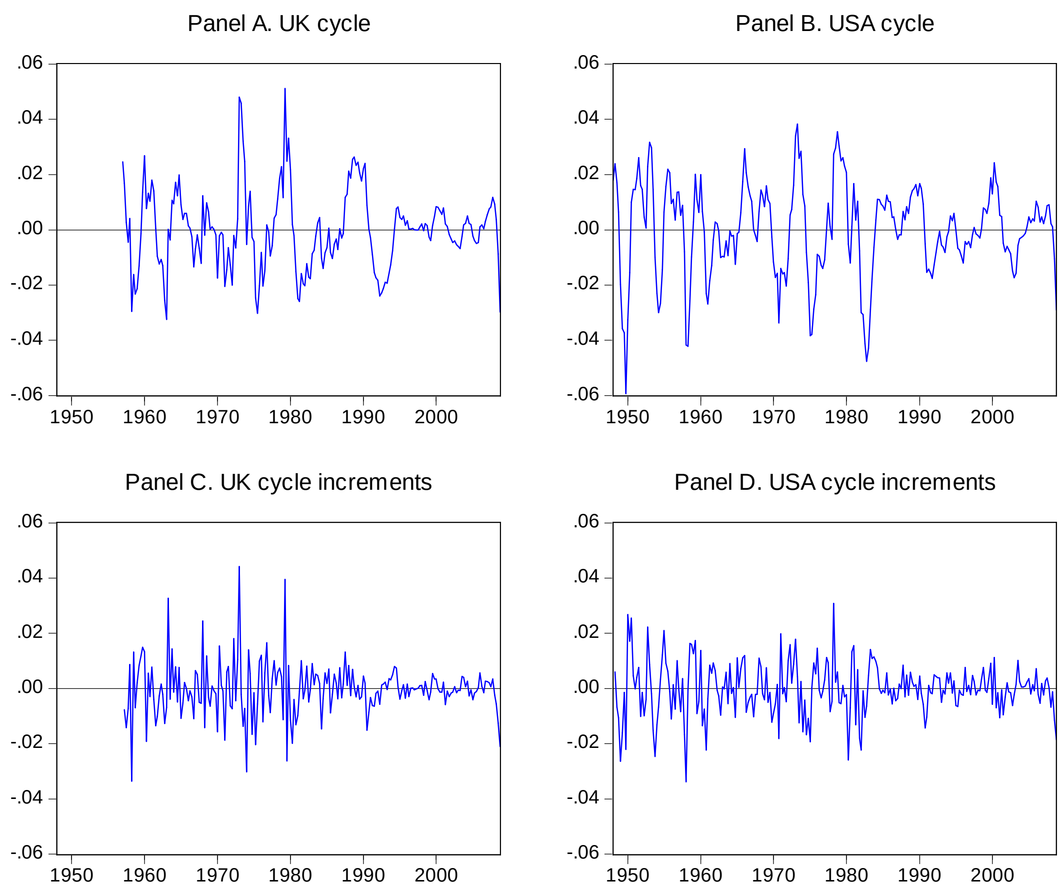

Figure 1: Panels A and B show UK and US cyclical components obtained by subtracting the trend computed with the Hodrick and Prescott filter with a smoothing parameter equal to 1600 from the logarithm of GDP, and panels C and D show their increments.

[1] However, the last quarter in the sample, as well as the first quarters in 2009, show a sharp decline in GDP in both countries. These movements would question some of these claims.

Table 1: Standard deviations

UK	1971:1 - 2008:4	1971:1 - 1995:4	1996:1 - 2008:4
$CYCLE$	0.015	0.017	0.006
$\Delta CYCLE$	0.009	0.011	0.004
UR (all)	2.500	2.740	0.930
UR (men)	2.940	3.360	1.200
UR (women)	2.090	2.060	0.640
ΔUR (all)	0.270	0.300	0.190
ΔUR (men)	0.340	0.370	0.240
ΔUR (women)	0.200	0.220	0.160
US	**1948:1 - 2008:4**	**1948:1 - 1987:4**	**1988:1 - 2008:4**
$CYCLE$	0.017	0.019	0.010
$\Delta CYCLE$	0.009	0.011	0.005
UR (all)	1.490	1.710	0.920
UR (men)	1.620	1.850	1.050
UR (women)	1.390	1.530	0.760
ΔUR (all)	0.390	0.450	0.240
ΔUR (men)	0.430	0.490	0.280
ΔUR (women)	0.370	0.430	0.210

Standard deviations of the variables indicated in the first column. *CYCLE* is the cyclical component obtained by subtracting the trend computed with the Hodrick and Prescott filter with a smoothing parameter equal to 1600 from the logarithm of GDP, $\Delta CYCLE$ is the increment in *CYCLE*, *UR* is the percentage unemployment rate and ΔUR is the increment in *UR*.

With regard to unemployment, the first important feature to analyze is whether the reduction in the amplitude of cyclical movements in GDP in the last few decades could have been propagated to unemployment. In order to analyze this possibility, Figure 2, panels A and B, shows the unemployment rates in the UK and USA. There also seems to be lower variability in the last few years, but this is not very clear. When taking increments in unemployment rates (see panels C and D in Figure 2), a similar pattern to that of cyclical GDP is observed more clearly: variability also decreases greatly in the last few years. This is confirmed in Table 1, where both unemployment rates and their changes record a lower standard deviation in the last sub-samples.

Table 2: Most extreme values

Ranking	UK $CYCLE$	US $CYCLE$	UK $\Delta CYCLE$	US $\Delta CYCLE$	UK ΔUR	US ΔUR
1	0.051 (1979:2)	-0.061 (1949:4)	0.044 (1973:1)	-0.034 (1958:1)	0.900 (1981:1)	1.6 (1975:1)
2	0.048 (1973:1)	-0.047 (1982:4)	0.039 (1979:2)	0.031 (1978:2)	0.900 (1980:4)	1.5 (1954:1)
3	0.045 (1973:2)	-0.042 (1983:1)	-0.034 (1958:2)	0.027 (1950:1)	0.800 (1980:3)	1.4 (1958:1)
4	0.033 (1979:4)	-0.042 (1958:2)	0.033 (1963:2)	-0.026 (1949:1)	0.700 (1981:2)	1.3 (1949:2)
5	0.031 (1973:3)	-0.041 (1958:1)	-0.030 (1974:1)	-0.026 (1980:2)	0.700 (1991:2)	1.1 (1958:2)
6	-0.030 (1975:3)	-0.041 (1982:3)	-0.026 (1973:3)	0.025 (1950:3)	0.5 (5 quarters)	1 (1953:4)
7	-0.029 (2008:4)	-0.040 (1949:3)	0.024 (1968:1)	-0.025 (1953:4)	-0.5 (5 quarters)	1 (1974:4)
8	0.026 (1988:4)	-0.039 (1949:2)	-0.021 (2008:4)	-0.022 (1982:1)		-1 (1950:3)
9	-0.025 (1981:2)	-0.038 (1975:1)	-0.020 (1975:2)	-0.022 (1960:4)		1 (1980:2)
10	0.025 (1988:3)	0.038 (1973:2)	-0.020 (1980:2)	0.022 (1952:4)		-0.9 (1958:4)

The entries are the most extreme cyclical components, increments in cyclical components and increments in unemployment rates with their dates in parentheses.

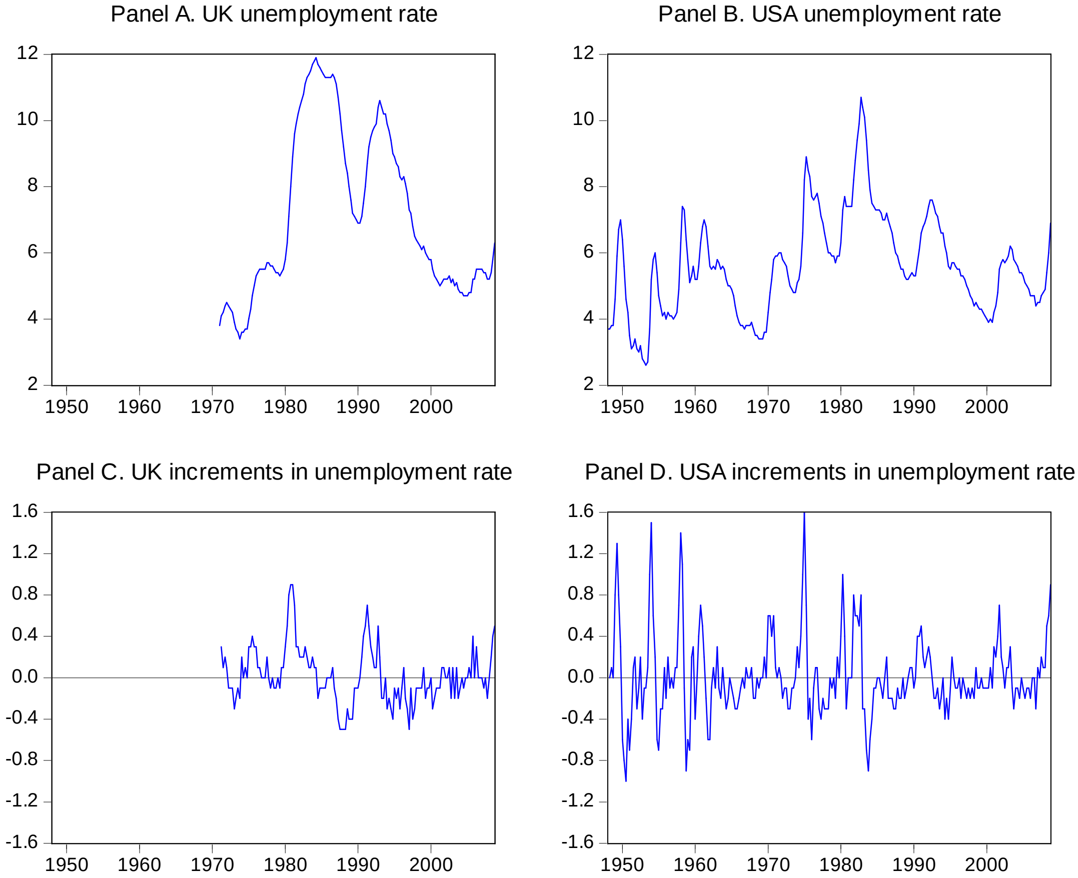

Figure 2: Panels A and B show UK and US percent unemployment rates, and panels C and D their increments.

Another interesting point reflected in Table 1 is that unemployment variables for women record a lower standard deviation than for men. For both countries, for the whole sample and for each sub-sample, women series vary less than their male counterparts and differences are often substantial. This could be at least partially due to female employment being more stable than male employment which would in turn be a possible consequence of the differences in gender employment by economic sector or activity. If sectors that are more sensitive to cyclical shocks have a large proportion of male employment, cyclical fluctuations will exert a stronger effect on male unemployment, and the relationship of cyclical movements with unemployment will be more intense for men than for women.

Finally, there seem to be some asymmetries in unemployment rates. Let us take US unemployment first. It is remarkable to note that in Figure 2, panel B, the unemployment rate in the US displays sharp peaks and much more rounded troughs: unemployment increases rapidly, but decreases more gradually. This point can be confirmed by looking at the increments in the unemployment rates in panel D. It is evident that the most extreme values are positive; in fact, eight out of the nine most extreme values are positive (see Table 2.) This can also be observed in the histogram of the empirical distribution of changes in unemployment rates (not shown here for reasons of space), skewed to the right with a skewness statistic equal to 1.04. This phenomenon may be not so evident for the UK in panel A, but it is also somewhat noticeable in the series of increments in the UK unemployment rate (see Figure 2, panel C); as shown in Table 2, the five most extreme variations in the UK unemployment rate are also positive. The empirical distribution is also skewed to the right with a skewness statistic of 0.73. Consequently, it seems reasonable to suspect the existence of asymmetries in unemployment rates and in their increments, both in the UK and the US.

3. Unemployment and business cycle

In order to analyze the relationship between unemployment and business cycle, different regression models could be specified. One of the first natural choices would be:

$$UR_t = \alpha + \beta \cdot CYCLE_t + u_t \tag{1}$$

i.e., the (quarterly) unemployment rate (UR) is regressed against the (quarterly) cyclical component of GDP ($CYCLE$). This first approach to the problem, however, does not take into account a long standing issue regarding empirical macroeconomics: the integration order of the involved variables. On the one hand, the cyclical component is, by construction, a stationary ($I(0)$) variable.[2] On the other hand, the unemployment rate requires further investigation since we

[2] In fact, the cyclical component obtained with the standard Hodrick-Prescott filter is always an $I(0)$ variable.

do not have beliefs a priori about its degree of persistence. Tables 3–4 present some unit root and stationarity tests for UR.[3]

Table 3: Unit Root Tests (Unemployment Rate)

	DF-GLS	Ng-Perron				Breitung
		MZ_α	My_t	MSB	MP_T	$T^{-1}\hat{\rho}_T$
UK	-1.246	-6.801*	-1.829*	0.269	3.653*	0.019
US	-0.997	-11.871**	-2.350**	0.198**	2.403**	0.014*

The entries are the Dickey-Fuller GLS detrended (DF-GLS), Ng-Perron and Breitung test statistics. In the case of the DF-GLS and Ng-Perron statistics, the optimal lags were automatically selected by using the Modified Akaike Information Criterion. The frequency zero spectrum was estimated by the AR-GLS detrended data method. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

Table 4: Stationarity Tests (Unemployment Rate)

	KPSS		Bierens-Guo			
	Bartlett kernel	QS kernel	$BG-1$	$BG-2$	$BG-3$	$BG-4$
UK	0.386*	2.625***	183.87***	158.76***	3.42	1.65
US	0.181	0.158	3.107	3.267	0.787	0.667

The entries are the Kwiatkovski et al. (KPSS) (1992) and Bierens-Guo (BG) (1993) test statistics. The optimal truncation lag for the KPSS was automatically selected by using Andrews' (1991) procedure. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

Regarding the UK unemployment rate, there is weak evidence against the unit root hypothesis and strong evidence against stationarity. There is however stronger evidence against the unit root null for the US unemployment rate and we have not found evidence against the stationarity hypothesis. Hence, we would conclude that the US unemployment rate is an $I(0)$ variable, whereas we have conflicting results for the UK unemployment rate. Thus, if we run the regression (1), we cannot preclude the possibility that, in the case of the UK, we could have an $I(1)$ variable regressed against an $I(0)$ variable, i.e., in Granger's(1995) terminology, we could have an unbalanced equation, in the sense that there could be an unwanted strong property on the left hand side of (1).

[3] We have computed the GLS-detrended augmented Dickey-Fuller (DF-GLS, 1996), the Ng-Perron (2001) and Breitung (2002) unit root tests and the Kwiaikovski et al. (1992) and Bierens-Guo (1993) stationarity tests.

In addition, abstracting from the aforementioned possibility of facing an unbalanced equation for the UK, the regression of (1) yields anomalous results. Simple regressions of the unemployment rate on a constant and current cyclical component or on a constant and single lag or lead show significant negative slopes, for both the UK and the US. Nevertheless, for both countries, multiple regression of the unemployment rate on a constant and current cyclical component, together with several leads and lags, shows high joint significance, but none of the individual coefficients are significant. Similar results are obtained when taking the increments in the unemployment rate. This phenomenon is due to the high collinearity existent among cyclical components in successive quarters and it is a direct consequence of the detrending procedure. Although the cyclical component is stationary, it displays a high autocorrelation, with first-order correlations of 0.76 and 0.83 for the UK and the US, respectively.

Among many other problems, this collinearity prevents studying the dynamics of the response of unemployment to cyclical conditions. In order to circumvent or mitigate these difficulties, first differences in the cyclical component were obtained. These differences register a much lower collinearity, with first-order correlations now equal to -0.12 and 0.27 for the UK and the US, respectively. Hence, variations in unemployment rates were regressed on a constant and changes in the cyclical component,

$$\Delta UR_t = \alpha + \sum_{i=-m}^{n} \beta_i \cdot \Delta CYCLE_{t+i} + u_t \qquad (2)$$

where ΔUR_t is the variation in the unemployment rate in quarter *t*, $\Delta CYCLE_t$ is the increment in the cyclical component in quarter *t* and u_t is the error term for the same quarter. Notice that with (2) we also solve the unbalanced equation problem we could face if the UK unemployment rate were a truly $I(1)$ variable, since differencing both sides of (1) renders stationary variables in any case.

Values for *m* and *n* (lag and lead selection) were chosen by using sequential *F* tests with a significance level set at 5%, and the conclusions were almost identical to those obtained using Schwarz's Bayesian Information Criterion. Table 5 shows the results obtained for the UK and the US.

Table 5: Regression results

	UK	US
$\hat{\alpha}$	0.017	0.0064
	(0.029)	(0.0153)
$\hat{\beta}_1$	-4.09**	-7.37***
	(2.00)	(1.72)
$\hat{\beta}_0$	-9.92***	-21.84***
	(2.30)	(1.75)
$\hat{\beta}_{-1}$	-10.58***	-14.19***
	(2.74)	(2.21)
$\hat{\beta}_{-2}$	-9.75***	-8.78***
	(2.56)	(1.39)
$\hat{\beta}_{-3}$	-8.90***	-4.70***
	(2.22)	(1.58)
$\hat{\beta}_{-4}$	-6.62***	-
	(1.85)	
$\hat{\beta}_{-5}$	-4.43**	-
	(1.75)	
R^2	45%	70%

Results of the regressions $\Delta UR_t = \alpha + \sum_{i=-m}^{n} \beta_i \cdot \Delta CYCLE_{t+i} + u_t$ for both the UK and the US, where ΔUR_t is the change in the unemployment rate in quarter *t* and $\Delta CYCLE_t$ is the change in the cyclical component in quarter *t*. *m* is equal to 5 and 3, respectively, for the UK and the US, and *n* is 1 for both countries. Values in parentheses are Newey-West robust standard errors. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

This table presents several important results: i) Excluding the constants, the regressors included are, as expected, always negative, thus indicating a negative relationship between variations in business cycle and changes in unemployment rates; ii) Excluding the intercepts, the regressors are always clearly significant; iii) Variations in unemployment rates depend not only on contemporaneous changes in cyclical conditions but also on cyclical changes that took place several quarters before (five quarters in the UK and three in the US) and, interestingly, on cyclical changes that will take place in the following quarter. Therefore, the labor market reacts immediately and with a delay of several quarters to cyclical shocks, but also anticipates imminent cyclical changes; iv) As expected, the values of the estimates corresponding to nearer lags are higher (in absolute values) than those corresponding to further lags. This indicates the labor market is more sensitive to the most recent cyclical conditions; v) As reflected by the coefficients of determination, dependence seems to be stronger in the US than in the UK, but endures longer in the UK than in the US.

It is also important to grasp the meaning of the estimates. In order to do so, let us simulate the effects of a hypothetical expansion on unemployment. Let us suppose that in a ‘typical’ expansionary quarter the cyclical component increases by one percentage point (sample standard deviations of increases in the cyclical component are 0.95% and 0.94% for the UK and the US, respectively.) Let us also consider an expansion comprising four consecutive expansionary quarters. According to the estimation corresponding, for example, to the UK, the change in the unemployment rate in the quarter before the beginning of the expansion will be $0.017 - 4.09 \times 0.01 = -0.0239$, that is, an approximate decrease of 0.02 percentage points. The change in the unemployment rate in the quarter when the expansion begins will be $0.017 - -(4.09 + 9.92) \times 0.01 = -0.1231$, that is, an approximate decrease of 0.12 percentage points. Table 6 shows quarterly and accumulated reductions in the unemployment rate in the different quarters for both countries. The unemployment rate decreases slightly in the quarter immediately preceding the beginning of the expansion. Then it strongly decreases over the four expansionary quarters. In the following quarters, the UK unemployment rate still decreases strongly, but the additional decrease in the US unemployment rate is weaker. The overall effect of these four successive expansionary quarters is remarkably similar in both countries: a decrease in the unemployment rate of 2.00 percentage points for the UK and a decrease of 2.22 points for the US. Of course, this is just a hypothetical exercise to examine the effects of a given expansion and this variation will not be permanent insofar as the cyclical component does not remain stable indefinitely.

Table 6: Predicted changes in unemployment

Quarter	UK		US	
	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$
-1	0.02	0.02	0.07	0.07
0	0.12	0.15	0.29	0.35
1	0.23	0.38	0.43	0.78
2	0.33	0.70	0.52	1.30
3	0.37	1.08	0.49	1.78
4	0.34	1.42	0.27	2.06
5	0.28	1.70	0.13	2.18
6	0.18	1.88	0.04	2.22
7	0.09	1.97	–	–
8	0.03	2.00	–	–
9	–	–	–	–

The entries are the expected reductions in the unemployment rate over a four-quarter expansion when the cyclical component increases by one percentage point each quarter, and zero elsewhere. The first column indicates the number of quarters since the beginning of the expansion; that is, expansion begins in quarter 0 and ends in quarter 3.

The results reported above show a clear relationship between business cycles and the unemployment rate. But this relationship could be different by gender. It has been long argued that labor markets may differ greatly by gender; participation rates, wages, sector employment shares, among many other aspects, could be different for men and women, and the influence of cyclical shocks on the unemployment rate could also differ. Additionally, there has been discussion over the decline or even disappearance of business cycles in recent decades. If business cycles are much weaker than they used to be, the influence of cyclical shocks on the unemployment rate could also have diminished significantly. In order to study these two possibilities, new regressions similar to (2) have been carried out but: i) using male and female unemployment rates instead of the total unemployment rate; ii) using two sub-samples approximately made up of the first two thirds and the last third of the total sample. Table 7 shows the results of these regressions and Tables 8 and 9 show the dynamic effect, under these new circumstances, of an expansion comprising four successive increases in the cyclical component equal to 0.01, for the UK and the US respectively.

Table 7: Additional regressions

	UK				US			
	Men	Women	1971-1995	1996-2008	Men	Women	1948-1987	1988-2008
$\hat{\alpha}$	0.025	0.005	0.051	-0.051	0.009	0.004	0.012	0.002
	(0.035)	(0.026)	(0.036)	(0.032)	(0.017)	(0.016)	(0.020)	(0.024)
$\hat{\beta}_1$	-6.45***				-7.55***	-5.83**	-7.62***	
	(2.37)				(1.94)	(2.43)	(1.87)	
$\hat{\beta}_0$	-12.80***	-4.68***	-9.73***	-21.52***	-25.69***	-15.87***	-22.22***	-21.02***
	(2.65)	(1.63)	(2.22)	(5.92)	(2.03)	(2.55)	(1.87)	(4.40)
$\hat{\beta}_{-1}$	-14.44***	-4.92**	-10.70***		-13.88***	-14.15***	-14.25***	-15.67***
	(3.40)	(2.17)	(2.79)		(2.29)	(3.09)	(2.39)	(3.65)
$\hat{\beta}_{-2}$	-12.59***	-5.76***	-10.50***		-9.93***	-7.63***	-8.49***	-11.38***
	(3.53)	(2.15)	(2.71)		(1.74)	(2.29)	(1.50)	(3.65)
$\hat{\beta}_{-3}$	-12.23***	-4.15***	-8.39***		-5.77***		-4.53***	-8.69**
	(2.54)	(1.59)	(2.17)		(1.86)		(1.68)	(4.27)
$\hat{\beta}_{-4}$	-8.17***	-3.67**	-6.78***					
	(2.16)	(1.51)	(1.68)					
$\hat{\beta}_{-5}$	-6.63***	-2.76**	-4.86***					
	(1.92)	(1.35)	(1.79)					
R^2	51%	16%	51%	23%	70%	52%	73%	53%

Results of the regressions $\Delta UR_t = \alpha + \sum_{i=-m}^{n} \beta_i \cdot \Delta CYCLE_{t+i} + u_t$ for both UK and US, where ΔUR_t is the change in the unemployment rate in quarter *t*, $\Delta CYCLE_t$ is the change in the cyclical component in quarter *t*, and *m* and *n* take different values in the different regressions. Values in parentheses are Newey-West robust standard errors.The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

Table 8: Predicted changes in unemployment (UK)

	Men		Women		1971-1995		1996-2008	
Quarter	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$
-1	0.04	0.04	-	-	-	-	-	-
0	0.17	0.21	0.04	0.04	0.05	0.05	0.27	0.27
1	0.31	0.52	0.09	0.13	0.15	0.20	0.27	0.53
2	0.44	0.96	0.15	0.28	0.26	0.46	0.27	0.80
3	0.50	1.45	0.19	0.47	0.34	0.80	0.27	1.06
4	0.45	1.90	0.18	0.65	0.31	1.11	-	-
5	0.37	2.27	0.16	0.81	0.25	1.37	-	-
6	0.25	2.52	0.10	0.91	0.15	1.52	-	-
7	0.12	2.64	0.06	0.97	0.07	1.58	-	-
8	0.04	2.68	0.02	0.99	-	-	-	-
9	-	-	-	-	-	-	-	-

The entries are the expected reductions in the UK unemployment rate over a four-quarter expansion when the cyclical component increases by one percentage point each quarter, and zero elsewhere. The first column indicates the number of quarters since the beginning of the expansion; that is, the expansion begins in quarter 0 and ends in quarter 3.

Table 9: Predicted changes in unemployment (US)

	Men		Women		1948-1987		1988-2008	
Quarter	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$	$-\Delta UR$	$-\Sigma\ \Delta UR$
-1	0.07	0.07	0.05	0.05	0.06	0.06	–	–
0	0.32	0.39	0.21	0.27	0.29	0.35	0.21	0.21
1	0.46	0.85	0.35	0.62	0.43	0.78	0.36	0.57
2	0.56	1.41	0.43	1.05	0.51	1.29	0.48	1.05
3	0.54	1.96	0.37	1.43	0.48	1.78	0.57	1.62
4	0.29	2.24	0.21	1.64	0.26	2.04	0.36	1.97
5	0.15	2.39	0.07	1.71	0.12	2.16	0.20	2.17
6	0.05	2.44	–	–	0.03	2.19	0.08	2.26
7	–	–	–	–	–	–	–	–
8	–	–	–	–	–	–	–	–
9	–	–	–	–	–	–	–	–

The entries are the expected reductions in the US unemployment rate over a four-quarter expansion when the cyclical component increases by one percentage point each quarter, and zero elsewhere. The first column indicates the number of quarters since the beginning of the expansion; that is, the expansion begins in quarter 0 and ends in quarter 3.

One main conclusion emerges from these tables. Male unemployment is more sensitive to cyclical conditions than female unemployment. This is the case in both the UK and the US, but the difference is much larger in the UK. An expansionary period comprising four successive increases of one percentage point in the cyclical component would lead to a decrease in the UK male unemployment rate of 2.68 percentage points, while the female unemployment rate would only decrease by 0.99 percentage points. The difference is much larger than in the US, where the men's unemployment rate would decrease by 2.44 percentage points, while women's rate would decrease by 1.71 points. With regard to changes over time in the response of unemployment to cyclical shocks, in the first sub-period UK unemployment was more responsive to cyclical conditions than in the second sub-period. The expansion comprising four consecutive expansionary quarters would have resulted in a decrease of 1.58 percentage points in the first sub-period and 1.06 points in the second. However, in the US the reductions are 2.00 and 2.22, respectively. Therefore, the change over time does not seem to occur in the US, where the sensitivity of the unemployment rate is very similar in both sub-periods.

One possible explanation for these facts could be the difference in employment by sector between men and women and over time. The ratio of employment in goods-providing sectors to employment in service-providing sectors is always much higher among men than among women, both in the UK and the USA. If unemployment in the first sectors were more sensitive to cyclical shocks than in the latter, one should expect male unemployment to be more sensitive to cyclical conditions than female unemployment, as turns out to be the case in both countries. The evolution of this ratio can also explain the different response from the UK unemployment rate over time. While the mean of this ratio is 59% in the first sub-period, it decreases to 33% in the second. This reduction may be associated to fact that the unemployment rate is less sensitive to cyclical conditions observed in the second sub-period. Nevertheless, in the US the mean of this ratio also decreases from 63% to 29% and no major changes are observed in the response of the unemployment rate to cyclical movements. Other factors, such as changes in labor legislation or dismissal procedures, probably play an important role over time.

4. Markov switching regime models

In the preceding section we investigated the link between unemployment and GDP by means of standard linear regression equations. However, such approach does not take into account the possible nonlinear relationships between the variables.

As seen in Section 2, unemployment series seem to be asymmetric. In what follows, first we formally test for unconditional symmetry of the unemployment rate series. In order to do so, we employ two recently proposed tests: the Bai and Ng (2005) and Racine et al. (2007) procedures, see Table 10.

Table 10: Unconditional symmetry tests

Series (UK)	$\hat{\pi}_3$	$\hat{\mu}_{35}$	$\hat{S}_\rho$
ΔUR (all)	1.507	2.305	0.000***
ΔUR (men)	1.652*	3.051	0.002***
ΔUR (women)	0.328	0.644	0.810
Series (US)			
ΔUR (all)	2.435**	7.478**	0.032**
ΔUR (men)	2.427**	6.724**	0.025**
ΔUR (women)	1.996**	5.099*	0.010**

$\hat{\pi}_3$ and $\hat{\mu}_{35}$ are the Bai-Ng (2005) test statistics, whereas $\hat{S}_\rho$ is the Racine et al. (2007) statistic p-value. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

The Bai-Ng test does not allow us to reject the null hypothesis of symmetry for any of the UK unemployment rate series (there is just a marginal 10% level rejection for the male series using the $\hat{\pi}_3$ statistic), whereas the Racine et al. test strongly rejects the null of symmetry for the total and male unemployment rate series. In addition, we find strong evidence against symmetry for all the US variables when using both testing procedures. These results could be explained by either asymmetric disturbances in the linear data generating processes, or possible nonlinear mechanisms driving the dynamics of the changes in the unemployment rates.

In order to delve deeper into this question, next we test for linearity using the Hong and Lee (2007) test. The results in Table 11 are rather conclusive. The null hypothesis of linearity cannot be rejected for any UK variable. However, it is clearly rejected for all the US unemployment rate changes series.

Table 11: Hong-Lee (2007) linearity test

Series (UK)	$\bar{p} = 10$	$\bar{p} = 20$	$\bar{p} = 30$	$\bar{p} = 40$	$\bar{p} = 50$
ΔUR (all)	0.114	0.090*	0.074*	0.069*	0.068*
ΔUR (men)	0.122	0.128	0.144	0.154	0.154
ΔUR (women)	0.486	0.334	0.211	0.150	0.114
Series (US)					
ΔUR (all)	0.038**	0.049**	0.058*	0.055*	0.046**
ΔUR (men)	0.000***	0.001***	0.003***	0.007***	0.010***
ΔUR (women)	0.009***	0.004***	0.002***	0.001***	0.001***

The entries are the p-values of the Hong-Lee test. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

Given the results of the symmetry and linearity tests, model (2) could be a plausible representation of the dynamics of the total and male UK unemployment rate changes series, whereas it is an inappropriate mechanism to explain the complex movements in US variables. Hence, alternatively, we estimate a set of Markov switching (MS) regime regression models.[4]

A stationary time series y_t is assumed to have been generated by an MS model with *M* regimes and *p* lags in the exogenous regressors if:

$$y_t = \alpha(s_t) + \sum_{k=0}^{p} \beta_k(s_t) x_{t-k} + u_t, \quad (3)$$

where $y_t = \Delta UR_t$, $x_t = \Delta CYCLE_t$, $u_t | s_t \sim N(0, \sigma^2)$, and the values of the intercept and the multipliers depend on the current regime represented by s_t. If we allow for regime-dependent heteroskedasticity, then we replace the assumption $u_t | s_t \sim N(0, \sigma^2)$ by $u_t | s_t \sim N\left(0, \sigma^2(s_t)\right)$.

This type of models was initially developed by Goldfeld and Quandt (1973), in response to economists' view of different behavior of variables during different

[4] The universe of available nonlinear models is large (TAR, STAR, ESTAR, LSTAR, etc.) Our goal however is not to fit the "best" nonlinear model, but rather to show how our conclusions can be enriched by alternative approaches.

cycle phases. The phase is represented by an unobservable state s_t, which takes a value of 1 (expansion) or 0 (contraction.)[5] The simplest specification is that s_t is the realization of a two-state Markov chain, where the probability of a change in regime depends on the past only through the value of the most recent regime:

$$P(s_t = j \mid s_{t-1} = i, s_{t-2} = k, \ldots, y_{t-1}, y_{t-2}, \ldots) = P(s_t = j \mid s_{t-1} = i) = p_{ij}.$$

The formulation of the problem, in which all parameters of interest are calculated as a by-product of an iterative algorithm similar in spirit to a Kalman filter, is due to Hamilton (1989.) The maximum likelihood estimates have been obtained by the EM algorithm. Results of the selected models using Schwarz's criterion are displayed in Tables 12 and 13.

Table 12: MS estimates (UK)

	ΔUR (all)		ΔUR (men)		ΔUR (women)	
	Expansion	Contraction	Expansion	Contraction	Expansion	Contraction
$\hat{\alpha}$	-0.11***	0.19***	-0.10***	0.23***	-0.09***	0.16***
	(0.02)	(0.02)	(0.03)	(0.03)	(0.02)	(0.02)
$\hat{\beta}_1$	-7.54**	-3.75**	-6.37*	-4.18**	-5.13*	-3.06*
	(3.47)	(1.57)	(3.28)	(1.81)	(2.80)	(1.58)
$\hat{\beta}_0$	-13.39***	-9.77**	-9.76***	-14.61***	-9.17***	-4.72***
	(2.90)	(1.65)	(2.90)	(1.89)	(2.79)	(1.66)
$\hat{\beta}_{-1}$	-9.94***	-11.25***	-8.98***	-19.27***	-7.51***	-5.20***
	(3.04)	(1.57)	(2.64)	(1.81)	(2.42)	(1.64)
$\hat{\beta}_{-2}$	-7.64**	-11.09***	-6.89**	-17.82***	-7.60**	-5.74***
	(2.83)	(1.58)	(2.76)	(1.82)	(2.64)	(1.56)
$\hat{\beta}_{-3}$	-7.93***	-7.77***	-7.41**	-12.98***	-4.43*	-4.43**
	(2.45)	(1.83)	(2.99)	(1.89)	(2.26)	(1.95)
$\hat{\beta}_{-4}$	-4.98*	-4.14**	-3.95	-4.31**	-4.87**	-2.47
	(2.93)	(1.72)	(3.15)	(1.76)	(2.45)	(1.76)
$\hat{\sigma}$	0.14	0.13	0.19	0.12	0.14	0.12

The entries are the estimated parameters for each regime. Values in parentheses are Newey-West robust standard errors. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

[5] As typically, we have chosen to specify a binary state variable, although we could represent a more flexible framework by allowing, for instance, for an additional intermediate state between pure expansion and pure contraction.

Table 13: MS estimates (US)

	ΔUR (all)		ΔUR (men)		ΔUR (women)	
	Expansion	Contraction	Expansion	Contraction	Expansion	Contraction
$\hat{\alpha}$	-0.08***	0.05*	-0.09***	0.05*	-0.03**	0.04
	(0.01)	(0.03)	(0.02)	(0.03)	(0.02)	(0.04)
$\hat{\beta}_1$	-6.33***	-3.51	-5.17**	-5.46**	-7.33***	2.51
	(1.79)	(2.35)	(2.31)	(2.63)	(2.26)	(3.22)
$\hat{\beta}_0$	-8.28***	-27.83***	-10.38***	-32.52***	-9.89***	-27.55***
	(1.83)	(2.45)	(2.13)	(2.69)	(2.17)	(3.27)
$\hat{\beta}_{-1}$	-1.28	-19.64***	-2.06	-19.25***	-2.57	-23.44***
	(1.87)	(2.35)	(2.39)	(2.63)	(2.47)	(3.23)
$\hat{\beta}_{-2}$	-2.84	-8.89***	-0.95	-12.83***	-8.68***	-5.04
	(1.92)	(2.38)	(2.21)	(2.74)	(2.31)	(3.25)
$\hat{\sigma}$	0.11	0.20	0.13	0.22	0.17	0.20

The entries are the estimated parameters for each regime. Values in parentheses are Newey-West robust standard errors. The superscripts *, ** and *** indicate significance at the 10%, 5% and 1% levels, respectively.

Concerning UK results, multiplier estimates for the total unemployment rate do not largely differ among regimes and standard deviation estimates are rather close. However, behind this result, a different fact is found for male and female unemployment. In absolute values, the estimates are somewhat larger in expansions than in contractions for female unemployment, while the opposite is true for male unemployment. In the case of the latter, the absolute estimates are much larger in contractions than in expansions. This marked difference between regimes implies deep asymmetry: the male unemployment rate is much more sensitive to contractions than to expansions. The same occurs for US unemployment variables and the differences between expansions and contractions are considerable in all cases. The symmetry test results shown in Table 10 are thus confirmed; absolute estimates are much higher in contractions than in expansions, implying US unemployment is much more sensitive to contractions than expansions.

In order to assess the effect of cyclical movements on the unemployment rate from a different perspective, simulations of the effects of hypothetical expansions and contractions on unemployment are also performed. Let us once again consider an expansion (contraction) period comprising four quarters with increases in the cyclical component equal to 1% (-1%.) Their effects on unemployment rates are shown in Table 14. Two main conclusions can be derived from this table. First, expansions diminish the unemployment rate to a smaller extent than contractions increase it in the case of both countries. In some cases the difference is very large; for example, an expansion period reduces the UK male unemployment rate by 2.67 percentage points, while a contraction period increases it by 4.99 points. Second, male unemployment in contractions is much more sensitive than female unemployment; for example, a contraction period in the UK increases the male unemployment rate by 4.99 percentage points, whereas the female unemployment rate only rises by 2.50 points.

Table 14: Predicted accumulated changes in unemployment (MS models)

	UK		US	
	Expansion	Contraction	Expansion	Contraction
ΔUR (all)	3.05	3.59	1.29	2.71
ΔUR (men)	2.67	4.99	1.35	3.14
ΔUR (women)	2.37	2.50	1.37	2.43

The entries are the expected accumulated reductions (increments) in the unemployment rate, up to the 7th (UK) and 5th (US) quarter, over a four-quarter expansion (contraction) when the cyclical component increases (decreases) by one percentage point each quarter and zero elsewhere.

Therefore, the simulations shown in Table 14 also seem to point to two important features in the relationship between unemployment and business cycles: the asymmetric relationship of unemployment with expansions and contractions and the differences between male and female unemployment over the business cycle. Though the evidence presented here seems to be fairly strong and robust, further research should be undertaken on these two points using different approaches.

5. Conclusions

The study of the relationship between unemployment and business cycles has received the attention of economic research for many decades. However, the literature is very sparse on important related topics such as the dynamics in-volved, stability over time, the existence of possible differences by gender and the possible asymmetric response of unemployment to cyclical movements. This scarcity is extreme when it comes to integrating these topics in a joint analysis.

In order to mitigate this situation, this paper examines the relationship between unemployment and cyclical conditions over the last few decades in two main economies: the United Kingdom, and the United States of America. A strong and definite dynamic relationship is found in both countries that lasts for several quarters and displays a reasonable and logical time profile. These findings allow us to simulate the effects of a hypothetical expansion comprising four successive expansionary quarters with a cyclical component equal to the typical value of 1%. The results obtained are similar for both economies; the expansion would decrease the unemployment rate by 2.0 and 2.2 percentage points in the UK and the US respectively. A closer look shows that there are clear differences by gender and over time. The effect of cyclical shocks is clearly stronger on male unemployment than on female unemployment, especially in the UK. Over time, the effect on total unemployment has decreased noticeably in the UK, maybe due to the growing importance of services or to the role of the public sector.

Numerous contributions have reported empirical evidence on unconditional asymmetry of labor variables in many countries. Preliminary evidence suggests that this is the case with the UK and US unemployment rates. Markov switching regime models are estimated to account for this feature. The estimated models effectively show an asymmetric response on behalf of the unemployment rate to cyclical movements. The unemployment rate usually reacts more strongly to negative cyclical shocks (contractions) than to positive ones (expansions), and this difference is particularly pronounced for male unemployment.

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