Business cycles, labour market entry and inequality

P. Cerqueira[*], L. Nhung[†], M. Portela[‡], S. Sousa[§]

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Abstract

The recent economic crisis has recalled the researchers' attention to the importance of external economic factors not only for the performance of labour markets but for the development of societies in general. The present paper aims at contributing to the knowledge of the impacts of the business cycle in Portugal, proposing to provide new insights on two main issues: the establishment of a relation between the business cycle and the evolution of income inequality and the impact of external conditions at the time of labour market entry on the individual's labour market performance. Stating by a brief review of the results found in the literature the first contribution of this article is to establish the Portuguese business cycle, based on OECD data, from 1970 until 2012. The relation between the business cycle and income inequality is then investigated, based on data from Quadros de Pessoal, motivating a subsequent analysis of the impact of entering the labour market in a boom versus in a recession, on individuals' labour market outcomes, suggesting that external economic conditions matter.

Keywords: Business cycle, inequality, LEED

JEL Classification: *C??*, C33, E32, J31

[*]University of Coimbra and GEMF

[†]University of Coimbra and Minho

[‡]University of Minho and NIPE

[§]University of Minho and NIPE. Corresponding author: ssousa@eeg.uminho.pt.

1 Introduction

The recent economic crisis has recalled the researchers' attention to the importance of external economic factors not only for the performance of labour markets but for the development of societies in general. The abrupt and persistent recession that several countries are still dealing with, translated into high unemployment rates, high job insecurity and low job mobility has posed tremendous challenges to policy makers, not only while addressing labour market phenomena but also some of their predicted social consequences such as inequality and poverty.

The establishment of the importance of the external economic factors, when searching for a strategy aiming at promoting economic growth and development, in a context of strong restrictions on the use of traditional economic policy instruments, is particularly crucial when the existent evidence on the impact of such factors is either scarce or not robust.

The aim of the present research twofold. Firstly, it aims to establish the impact of the business cycle on income inequality in Portugal, following a vast strand of literature on this issue, and secondly to shed some light over how the business cycle may lead to such income inequality. The idea here is to investigate whether there is a consistent relation between the business cycle and the evolution of income inequality in Portugal and to what extent this relation is robust and, related to this objective but not necessarily dependent on the results found, to investigate whether the business cycle matters to specific labour market outcomes such as the wage level and the wage persistence of the individuals.

Stating by a brief review of the results found in the literature, both on the relation between the business cycle and inequality and on the impact of external conditions at the time of labour market entry on wages (level and persistence) aiming at providing context for future results, the first contribution of this article is to establish the Portuguese business cycle. Dating the Portuguese business cycle relies upon a thorough exercise from a methodological perspective, based on OECD data, which allows for the identification of the peaks, recession and growth periods, from 1970 until 2012. Built on the results provided by this initial exercise, the relation between the business cycle and inequality follows, based on data from Quadros de Pessoal, a LEED covering the Portuguese private sector, over a period of around two decades. The results suggest that a deeper analysis is needed in order to identify and better understand the impact of external labour market conditions on labour market outcomes, motivating the subsequent analysis of the impact on entering the labour market in a boom versus a recession, taking the years of 1992 (boom) and 1993 (recession) as entry dates and following the workers for around two decades. The results suggest that external labour market conditions matter in terms of wage level and persistence and job mobility and that individual characteristics (e.g. education) should not be ignored when determining such impacts.

2 Business cycles and inequality

It is widely known that business cycle has a substantial influence on income disparities. Nevertheless, both empirical and theoretical studies fail to reach a consensus on the sign of this effect. In general, many researches found that the trend of earnings inequality over economic fluctuations is pro-cyclical, some others stated that it is mixed, or countercyclical while the others concluded that it depends .

Early work on inequality evolution by Mendershausen (1946) and Kuznets (1953) reported that US inequality follows an anti-cyclical pattern in the inter-war years whereas the trend is mixed during the postwar periods. Kuznets (1955), later, as tracking business cycle using time series GDP data, explained the latter finding by hypothesizing that the relationship between economic growth and income inequality has a inverted U-shape, which is the curve named after Kuznets. Inequality first rises, peaks and later falls as the economy grows. A number of subsequent researches is consistent with Kuznets' conjecture such as Blinder and Esaki (1978), Nolan (1988) or Robinson (1991). Barron (2000), focusing on the 3 stage least square method over the cross-section country data, also added that the relations between inequality and GDP growth is negative in the initial stage of the economy and the trend changes the sign at the certain level of its development. However, this finding is not uniformly reported. Some other authors instead suggested that the relation should be negative (Forbes (2000); Simões, Andrade et al. (2013))

On the other hand, Li, Squire et al. (1998) argued that the Kuznets curve fits better with the cross-section of country at a point in time than the series of data within countries, implying that this pattern of inequality is more likely to stem from historical difference among countries than from the development progression of countries themselves. Additional, the validity of Kuznets curve theory is violated by many historical evidences observed from development economics such as the « East Asia miracle » or the « autocratic disaster » experiences (Robinson and Acemoglu (2002)). Instead, Hoover, Giedeman et al. (2009), by using the threshold and momentum models of co-integration analysing consistent time series data for one country, proposed that inequality has an asymmetric adjustment over the business cycle, in which economic downturns caused by an increase in unemployment raises the income inequality whereas negative shocks to unemployment just short-lived affect in this dispersion. Nevertheless, GDP, which is used to represent for business cycle, was found to be not co-integrated with the distribution of income in the long-run.

Alternatively, Barlevy, Tsiddon et al. (2004) by taking into account the drive of technological changes argued that recession is likely to amplify the long-run trend, i.e during the recession periods, inequality increases more rapidly when its long-run trend is increasing and more rapidly decreasing if the trend is down-warding. They pointed out that, in time of falling inequality, the less workers has less incentives to take advantages of downturns to catch up to those who are master already in the new technology and vice versa in the period of increasing inequality. Moreover, their hypothesis is consistent with the US history in the middle of the last century.

Though there indeed exists different patterns of inequality over business cycle with respect to gender, not many studies are found to pay attention on examining this discrepancy. Machado and Mata (2001) stated that men are more likely to contribute to the rise in wage inequality. They exhibited that the income distribution of women is to the left of men's and the gender gap is bigger for high-paid jobs; men's income distribution is more dispersed than women's, as a result. In addition, Bonhomme and Hospido (2012) examined the cycle of earning inequality in Spain from 1988 to 2010 and concluded that there exists the different inequality trend over business cycle between female and male employees in which male earning inequality consistently follows a countercyclical evolution while the female inequality tends to pro-cyclical in some periods and countercyclical in the others.

The positive correlation between average inflation and income inequality is found by many studies such as Al-Marhubi (1997), Romer and Romer (1999) or Albanesi (2007). Albanesi (2007) explained that this correlation is the outcome of distributional conflicts underlying fiscal policy, as observing cross-country evidences from 1966 to 1990. Mainly relying on the hypothesis that low income households are more vulnerable to inflation and government expenditure is exogenous, he explained that in countries where government spending is financed in favour of monetary expansion, increasing inflation makes low income households lost more than high income ones since even though it's more costly to hold money, low income households has to cash for a greater fraction of their purchases in comparison with high income ones (Erosa and Ventura (2000)).

Acemoglu (1999) proposed that schooling and earnings inequality are positively correlated. By theoretical approach, he stated that the economy tends to switch from a pooling to a separating labor demand equilibrium , in which higher-quality jobs are targeted at the skilled and low-capital jobs are allocated for the unskilled as the substantially increasing proportion of college graduates leads to the raising productivity gap between unskilled and skilled workers over time. As a result, the earnings gap between the low-skilled and high-skilled increases allowing the up-warding wage dispersion trend. This proposal was also empirically proved by Chinhui, Kevin et al. (1993), Deschênes (2001) or Autor, Katz et al. (2008). Deschênes (2001) added that wages are an increasingly convex function of years of education ; for instant, the wage gap between college post-graduates and college graduates is likely to raise more than its between college graduates and high school graduates.

Several studies, including Card (1992), Machin and Manning (1994), Autor, Katz et al. (2008) or Centeno and Novo (2009), concluded that the influence of minimum wages on earnings inequality is negligible. Autor, Katz et al. (2008) explained that since minimum wage has little correlation with upper half of earnings distribution and the large increase in the upper-tail of earning distribution takes responsibility for the rise in inequality in the US over the last 2 decades, it is less likely that a change in minimum wage could cause a big change for the upper-tail earnings and inequality. However, other authors, such as Card and DiNardo (2002), Lemieux (2006) and Bárány (2011), argued that lower minimum wages play an important role in the increase of inequality in the US since 1980. The main reason that he suggested is minimum wage affects inequality through three channels. Particularly, a reduction in minimum wages widens the ranges of worker abilities, reduces the role of education by making it easier to find jobs , hence it increases the difference between any two percentiles of wage distribution and affects the inequality at not only the bottom half but the top end of wage distribution.

Unemployment variable plays an important role in explaining the dispersion of income. According to González and Menende (2000), 43% of the total increase in inequality can be explained by the increase in unemployment. However, their correlation is still ambiguous. Numerous studies consistently suggested that there is of positive relation between unemployment and inequality, in which a fall in income resulted from the rise in unemployment leads to a higher inequality. Since unemployment tends to affect more the less skilled and the low-paid, an unemployment shock decreases much more the earnings of the bottom half of the distribution and then widen the wage gap between groups of earnings distribution (Mocan (1999), Barlevy, Tsiddon et al. (2004)). However, time series evidence doesn't always support the notion of a positive relationship between unemployment and wage dispersion . For example, the reduction of the UK unemployment rate was notable in the second half of the 1980's and after the economic crisis of the early 1990's whereas its inequality rose at a very fast pace between 1979 and 1991. Gregg and Machin (1994) relied on the reason of breaking structure to explain for that observation. Martínez, Ayala et al. (2001) examining the household income instead of individual income concluded that the correlation is not necessarily positive since family income typically comprises incomes from several employees, thus the loss in earnings of a member could be offset by the gain of the others.

3 Labour market entry, wage levels and persistence

There is a rich strand of literature linking individuals' outcomes in the labour market, with the economic conditions at the time of entry. However, this link is not consensual depending on the theoretical framework that supports its analysis. Labor demand shocks or recessions will have a limited effect on wages and career paths, according to neo-classical models, being leisure preferences and human capital investments the main determinants of labour market outcomes. Instead, other theories, based on segmented labour markets or, more recently, on job search, contacting and efficiency wages will tend to attribute a role to labour market shocks on individuals' outcomes. Nevertheless, the nature of their effect may still be the object of controversy (e.g. temporary vs persistent vs permanent) (Oreopoulos et al, 2006). When other aspects of the labour market are being considered, such as job mobility, wage differentials and job attachment, the results found in the literature have been able to establish that economic shocks play an important role (Oreopoulos et al, 2006).

A common approach to the analysis of the effect of the business cycle at labour market entry is to investigate the impact of graduating in a recession (Oreopoulos et al, 2006, Genda et al, 2010, Kahn, 2010, Yu et al, 2014) either on employment and earnings, employment quality or entrepreneurship. The results suggest that there is evidence of large, negative and persistent consequences and that the dimension of the effect may vary between countries and the skill level of the individual. Moreover, there is a negative impact on entrepreneurship, translated in a smaller probability of starting a business, although the fraction of self-employed tend to rise (Yu et al, 2014).

Despite this growing evidence on the importance of external economic conditions to labour market performance, a dispute on their relative importance compared to internal labour market conditions prevails (Galuscak et al, 2012).

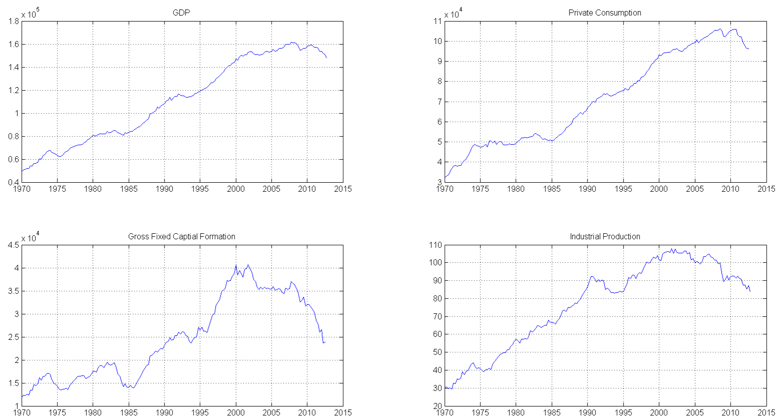
4 Dating the Portuguese business cycle

This section dates the Portuguese business cycle and marks the recession periods. Instead of using only one methodology, we used different methodologies (both multivariate and univariate) and compare the results in order to get a more robust and clear picture The methodologies used are the multivariate Markov Switching VAR proposed by Krolzig (1997), and two univariate methodologies, one proposed by Harding and Pagan(2002) based on the algorithm of Bry and Boschan(1971) and the other proposed by Artis et all. (2004). The following subsections will in turn describe the data used, give a brief description of each methodology, present the results obtained and the last one compares the results of the different methods and marks the expansion and recession periods.

4.1 A brief description of the data

We have collected quarterly seasonally adjusted data for the GDP, the Private Consumption, the Gross Fixed Capital Formation and the Industrial Production in real terms since the first quarter of 1970 until the third quarter of 2012[1] from the OECD database. Figure (1) shows the evolution of each one of these aggregates.

Figure 1: Evolution of main economic aggregates - Portugal - 1970q1 - 2012q3



This figure shows that all the aggregates had a steady increase since 1970 until 2000 with some short-lived periods of stagnation and/or recession. The first period of decline was around 1975, when Portugal was hit by two shocks: the first oil-crisis and the 1974 revolution that overturned a dictatorial regime and restored democracy. These years were characterized by a reduction on GDP, GFCF and Industrial Production, however Private Consumption was stable. Afterwards, the aggregates which had suffered a decrease experienced a fast recovery and kept growing until 1983. On the contrary, the Consumption was stagnated until 1980 and from 1980 until 1983 the increase was just marginal. In 1983, Portugal was at the verge of bankruptcy and called for financial assistance to the IMF[2]. The reforms and austerity measures taken at that time led to a reduction of the GDP, Private Consumption and GFCF, but curiously not of Industrial Production. After the end of this assistance program, Portugal joined the EEC (later

[1]Henceforward we will designate the periods as YYYYqX. For instance 1982q2, means the second quarter of 1982.

[2]Note that since 1970 period Portugal called three times for financial assistance. The first two financial assistance were provided by the IMF, as the last one was provided by the IMF together with the ECB and EU. In the first time a stand-by arrangement was concluded in May 1978 covering a year period. The second time the arrangement was concluded in September 1983 covering the period from October 7, 1983 to February 28, 1985 and the last one was set in May of 2011 for a period of three years.

EU) at the 1st of January of 1986 and experienced almost a full decade of high growth rates on all aggregates. This growth was briefly interrupted around 1992/1993 (with the exception of Private Consumption that continued to rise) and proceeded until 2000.

After 2000, the picture is grimmer. Industrial Production and GFCF started a decade long term of decline, while GDP fuelled by Private Consumption continued to raise until 2007. In 2007, Portugal, also felt the global impact of the sub-prime crisis, when all the aggregates declined (being that this decline was specially pronounced in GFCF and Industrial Production). The government tried to counteract this crisis with a stimulus program visible in the growth of GDP and Private Consumption, unfortunately GFCF and industrial Production stagnated or even declined. On May 2011, the imbalances of a full decade of GDP growth fuelled by private consumption and the sub-prime crisis led Portugal to ask, once more, for international financial assistance. This time the assistance was provided by the IMF, the ECB and the EU (the so called "Troika"). The reforms and austerity plans put in place at that time led to a mild decline in Industrial Production, but to an high decline in GDP, Private Consumption and due to financial restrictions to a major decline in Investment. We should note that while GDP and Consumption returned to the levels observed around 2000, the GFCF returned to the levels of 1995.

From this short analysis seems that since 1970 until 2013 we can point out four or five recessions: the first around 1975, the second around 1983-1985, the third around 1992 and, after 2000, we can say that in 2001 and in 2008 we can suspect that these two years were recession periods, and after 2011 until now we had, definitely, a major recession. The doubt is if 2001 is really a recession or just a stagnation and if 2008 recession can really be separated from the one of 2011 onwards. The following two subsections will try to measure more accurately the dating of the recessions, namely in which quarter has the economy peaked and in which quarter has it reached a trough.

4.2 The univariate methodologies

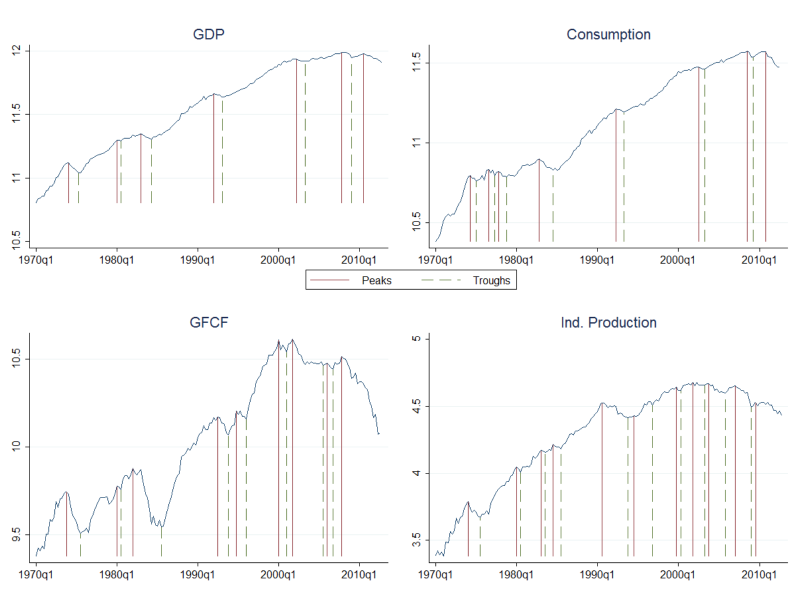
In this subsection we present the results of business cycle dating using two methodologies. The first is the algorithm associated with the NBER and set out in Bry and Boschan (1971) for monthly observations and adapted to quarterly series by Harding and Pagan(2002) and known as the BBQ. The second methodology was proposed by Artis et all. (2004) and is based on the theory of Markov Chains (MC) that enforces alternation of peaks and troughs and duration constraints giving the probability of shifts. The following points will give a brief description of each methodology and present the results to Portugal between 1970q1 and 2012q3.

4.2.1 The BBQ methodology

The original Bry and Boschan(1971) methodology includes a number of steps, but at the heart of the second step[3] is a definition of a local peak (trough) as occurring at time t whenever $y_t > (<) y_{t \pm k}$ for $k = 1, ..., K$ with K usually set to five. The main criteria relating to the this step is that a phase must last at least 6 months and a complete cycle should have a minimum duration of 15 months. Harding and Pagan (2002) have adapted the previous methodology to quarter data, calling it the BBQ algorithm. They set the K to 2 such that a phase lasts at least 2 quarters and a complete cycle has to last at least 5 quarters.

We applied the BBQ algorithm to the logarithm of the four main economic aggregates described in the previous section. Figure (2) and Table (1) show the estimated peaks and troughs for each one of the economic variables.

Figure 2: Turning points using the BBQ algorithm - 1970q1 -2012q3



From these estimates if we use only the GDP to classify the periods we get as recessions: 1974q1-75q2, 1980q1-80q3, 1983q1-84q2, 1992q1-93q1, 2002q2-03q2 , 2007q4-09q1, 2010q3-sample end. If we use a more broad picture and consider that a crisis happens when the majority of the variables between Consumption, GFCF and Industrial Production are in a recession we get a slighter different dating. The major differences are that some crisis are longer: the one in the mid-eighties spans now from 1982q4 to 1985q3; the 2002/03 crisis spans from 2001q4 to 2005q3 and the last two are merged into one single recession period from 2007q4 to the sample end. It also identifies new

[3]We should note the Bry and Bosch propose before to identify peaks and troughs, to identify outliers and replace them by suitable values.

Table 1: Peaks and Troughs- Portugal - 1970q1-2012q3

GDP		Consumption		GFCF		Ind. Production	
Peak	Trough	Peak	Trough	Peak	Trough	Peak	Trough
							1971q1
1974q1	1975q2	1974q2	1975q1	1973q4	1975q3	1974q1	1975q3
1980q1	1980q3	1976q3	1977q2	1980q1	1980q3	1980q1	1980q3
1983q1	1984q2	1977q4	1978q4	1982q1	1985q3	1983q1	1983q3
1992q1	1993q1	1982q4	1984q3	1992q3	1993q4	1984q3	1985q3
2002q2	2003q2	1992q2	1993q2	1994q4	1996q1	1990q3	1993q4
2007q4	2009q1	2002q3	2003q2	2000q1	2001q1	1994q3	1996q4
2010q3		2008q3	2009q2	2001q4	2005q3	1999q4	2000q2
		2010q4		2006q1	2006q4	2001q4	2003q2
				2007q4		2003q4	2005q4
						2007q1	2009q1
						2009q3	

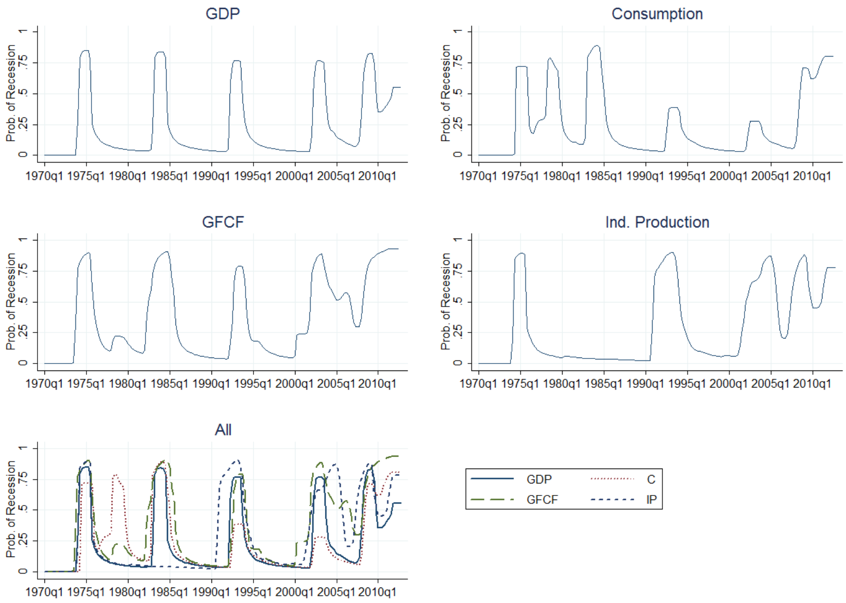
recession periods in 1994q4-96q4 and 2000q1-00q2,

4.2.2 The Markov Chains based methodology

The Markov-switching autoregressive time series model has become increasingly popular since Hamilton's (1989) application of this technique to measure the US business cycle. In this paper we will use a dating algorithm proposed by Artis et all. (2004) which is based on an appropriately defined Markov chain that enforces alternation of peaks and troughs and duration constraints concerning the phases and the full cycle in line with the duration and phase constraints imposed in the BBQ methodology. This methodology is more informative than the BBQ because instead of defining a period as being part of a expansion or a recession it estimates probability of being in either of these states. If the estimated probability is 100% for one state the conclusion is obvious, but if the estimated probability is around 50% it can be more tricky to access if it is a expansion or a recession. Appendix A gives a description of the methodology.

Analysing the results (see Figure(3)) for each series until 2000 we identify three major periods with high probabilities of recession: the first one is in 1974-75 when GDP has a higher probability of being in a contractionary period from 1974q2 until 1975q3. Consumption lags one period, as GFCF and Industrial Production starts one period earlier, but finishes almost at the same time as the recession marked by the GDP. We should also refer that the consumption was in a recession period from 1978q2 until 1979q3, but not the other variables. So we shouldn't consider this consumption adjustment as an overall crisis. The second crisis period is in 1983-84, as GDP is in recession from 1983q2 until 1984q3, the Consumption contractionary period starts at the same time but finishes 2 quarters latter and GFCF leads the crisis by 3 quarters

Figure 3: Recession Probabilities



and only returns to an expansionary phase in the end if 1985. It is also important to mention that in this period, the Industrial Production is the only series that does not show a recession period. Finally, the third crisis period was 1992-93 as marked by the GDP (from 1992q3 until 1993q3), GFCF recession period started almost one year earlier (1992q4) and finishes two quarters after the GDP contraction. As for the Industrial Production the contraction period starts in the beginning of 1991q1 and continues until 1994q1. In this period the recession probability of Consumption is in some periods almost 40%, so it is difficult to say if at that time it was in expansion or in a recession.

After 2000 the recession probabilities show a more complex picture: GDP was in a recession period from 2001q2 until 2003q3 and then again from 2008q2 until 2009q3 and from 2011 until the sample end. As for GFCF and the Industrial Production the crisis went from early 2002 until 2006q3 and 2005q3 respectively, while consumption kept very low levels of recession probabilities. It seems that the short GDP recovery in 2004 and 2005 was foremost led by internal demand as the investment and production continued to decline. After 2008q2 (and even earlier in the case of industrial production: 2007q3) until the sample end, all the variables point to Portugal to be in a recession.

4.3 The MS-VAR model

First attempts at the analysis of international business cycles with multivariate Markov-switching models have been undertaken by Phillips (1991), Filardo and Gordon (1994) and Krolzig (1997). In this paper we follow the approach proposed in Krolzig (1997) as the first two are restricted to a two regime model. The advantage of this methodology, henceforward MS-VAR, over the previous methods is that we can combine the information in different variables to characterize the different states, and so, we can identify more than just 2 states as the different variables might show different behaviours in the different recessions or/and expansions.

A MS-VAR can be seen as a generalization of the linear VAR. So consider p^{th} order autorregression for the K-dimensional time series vector y_t :

$$y_t = v + A_1 y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + u_t$$

if $u_t \sim N(0, \Sigma)$ and $A(L) = I_K - A_1 L - ... - A_p L^P$ as the $K \times K$ dimensional lag polynomial without roots inside the unit circle, the previous equation is known as the intercept form of a stable Gaussian VAR(p) model which can be rewritten as:

$$y_t - \mu = A_1(y_{t-1} - \mu) + A_2(y_{t-2} - \mu) + ... + A_p(y_{t-p} - \mu) + u_t$$

where $\mu = \left(I_k - \sum_{j=1}^{p} A_j \right)^{-1} v$ is the $(K \times 1\ )$ dimensional mean of y_t.

If the time series are subject to regime shifts, then this model might be inappropriate and so a MS-VAR can be considered as a regime switch framework. The general idea is that the parameters of the data generated process of y_t are dependent of the (non-observable) regime switch variable s_t, which represents the probability of being in a different state of the world. In this case the MS-VAR can be written as:

$$y_t - \mu(s_t) = A_1(s_t)(y_{t-1} - \mu(s_t)) + A_2(s_t)(y_{t-2} - \mu(s_t)) + ... + A_p(s_t)(y_{t-p} - \mu(s_t)) + u_t$$

where u_t ~$N(0, \Sigma(s_t))$ where $\mu(s_t), A_1(s_t), A_2(s_t), ..., A_p(s_t)$ and $\Sigma(s_t)$ are parameter shift functions describing the dependence of the parameters on the realized regime $s_t \in \{1, 2, ..., M\}$. Alternatively, we can consider that the regime shift is not in the mean (which implies a one time jump at the time of the regime shift), but in the intercept (implying a smooth transition between regimes). In that case the MS-VAR is written as:

$$y_t = v(s_t) + A_1(s_t) y_{t-1} + A_2(s_t) y_{t-2} + ... + A_p(s_t) y_{t-p} + u_t$$

Finally we need a characterization of the regime shift. The special characteristic of the Markov Switch models is the assumption that the unobservable realization of the regime s_t is governed by a discrete time Markov stochastic process,which is defined by the transition probabilities:

$$p_{ij} = \Pr(s_{t+1} = j | s_t = i) \quad , \quad \sum_{j=1}^{M} p_{ij} = 1$$

More precisely, it is assumed that s_t follows an irreducible ergodic M state Markov process with the transition matrix P formed by the p_{ij} elements.

Anyway, there's no need to have all the model parameters regime dependent, and in empirical applications can be more useful to fix some of them across regimes. Once we define the appropriate model we can estimate it by Maximum likelihood (ML) methods based on an implementation of the Expectation Maximization (EM) algorithm proposed by Hamilton (1990) for this class of models[4].

In this paper we applied the methodology to two sets of variables: {GDP, Consumption; GFCF} and {Ind. Production, Consumption; GFCF}, considering that the regime switching was done over the mean of each series and experimented with and without variance changes. However, when we included regime switches in the variance, the estimated variances were similar across regimes and there was no gain over the models without variance regime dependence. Therefore, due parsimony, we opted to keep just the mean regime dependent. Also, as the model has the assumption that the variables have to be stationary, we tested for the presence of unit roots. All of them were I(1) and so we applied the model to the logged first differences[5].

The next step was to decide the number of lags and states. Table (2) shows the results for the Akaike (AIC), the Schwarz (BIC) and the Hannan-Quinn (HQ) information criteria[6]:

Regarding the first set of variables {GDP, Consumption and GFCF} the minimum of the BIC selects 5 states and 1 lags, i.e., a MS-VAR(5,1) while the AIC and the HQ select a MS-VAR(4,2).

We estimated both models. Table(3) shows the estimated mean for each variable in each one of the states and figures (4) and (5) show the probability of the economy to be in each state.

Looking, first, to the MS-VAR(4,2) model, it is clear that state 1 marks the recession periods, when all variables had chain growth rates between -1,2% and -4,5%. The periods when the probability is well above 50% are: 1974q4-1975q2; 1989q2; 1991q1 and 2009q1. Although it captures some periods of the major crisis estimated by the

[4]For more details on the estimation methodology see Krolzig(1997)

[5]Results available from the authors.

[6]We tested for a maximum of 5 states and 4 lags with 2 or 3 states and 3 lags for 4 or 5 states.

Table 2: Information Criterias to decide on the number of lags and states

Variable sets		{GDP, Consumption; GFCF}			{Ind. Prod., Consumption; GFCF}		
States	Lags	AIC	BIC	HQ	AIC	BIC	HQ
2	1	1777.92121	1824.32747	1796.76164	2067.03808	2113.44434	2085.87851
	2	1752.34611	1826.59611	1782.4908	2056.82822	2131.07823	2086.97291
	3	1752.03779	1854.13155	1793.48674	2062.19131	2164.28507	2103.64026
	4	1760.01927	1889.95678	1812.77247	2067.0505	2196.98801	2119.80371
3	1	1751.14215	1806.82965	1773.75066	2050.37201	2106.05952	2072.98053
	2	1741.25221	1824.78347	1775.16499	2033.40744	2116.9387	2067.32022
	3	1739.90048	1851.27548	1785.11751	2026.14502	2137.52003	2071.36205
	4	1751.2506	1890.46935	1807.77189	2023.19651	2162.41527	2079.7178
4	1	1733.99484	1798.96359	1760.37144	2034.577	*2099.54576	2060.95361
	2	*1711.05803	1803.87053	*1748.73889	*2017.90867	2110.72118	*2055.58953
	3	1735.97953	1856.63579	1784.96465	2021.79622	2142.45248	2070.78134
5	1	1719.93991	*1794.18992	1750.0846	2038.51506	2112.76507	2068.65975
	2	1718.90814	1821.0019	1760.35709	2019.86705	2121.9608	2061.31599
	3	1714.87169	1844.8092	1767.6249	2025.9446	2155.88211	2078.69781

Table 3: Estimated mean over each Economy State

State	GDP	Consumption	GFCF	GDP	Consumption	GFCF
		MS-VAR(4,2)			MS-VAR(5,1)	
1	-2.05524	-1.20583	-4.53853	-0.85438	-0.72779	-3.60182
2	0.54607	0.60161	0.39325	0.75913	0.76673	0.99315
3	1.05523	-0.38348	0.31001	1.94775	-2.42232	2.79925
4	2.46786	3.57429	1.77515	2.02788	2.66652	1.54635
5				3.43609	5.26755	5.24235

Figure 4: MS-VAR(4,2) States Probabilities

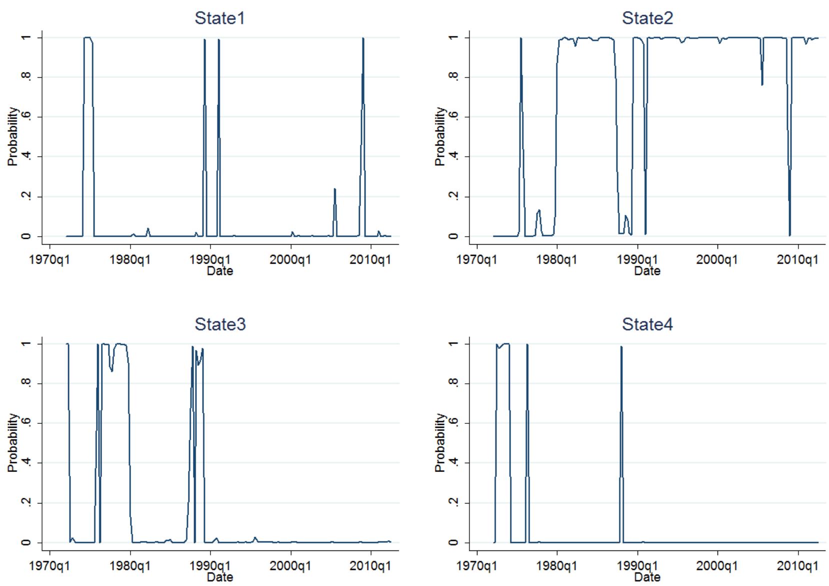
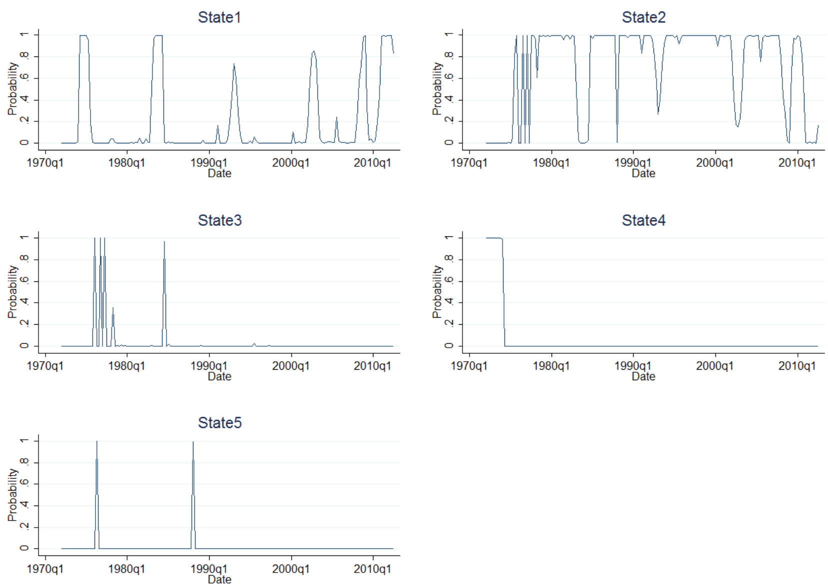


Figure 5: MS-VAR(5,1) States Probabilities



previous models, the estimated means imply yearly recession rates of 5% to 18%. so it seems that smaller recessions are not captured in this regime but mingled in the others.

As for the MS-VAR(5,1) it is, also, clear that recession periods are described by state 1 which comprises with high probabilities the following periods: 1974q2-1975q2, 1983q1-1984q2, 1993q1-1993q2, 2002q2-2003q1, 2008q2-2009q2 and 2011q1 to the sample end. As for the other 4 states, state 2 seems to be the normal expansion periods, while state 3 captures periods where although the economy is growing, consumption declines. State 4 captures the period previous to the democratic regime, while state 5 captures two historical marks: the re-normalization of the economy after the democratic revolution of 1974 and the political and economic turmoil that lasted the following two years (1976q2) and the admission of Portugal to the EEC (later, EU) in 1986 and the first single party majority government in 1987 that lead to a period of political certainty (1988q1).

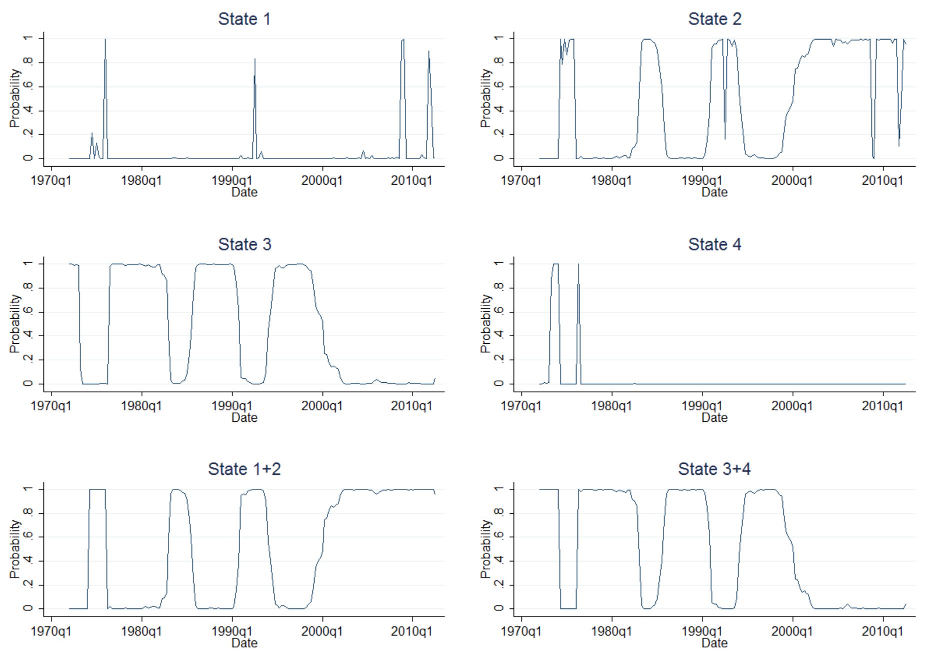
If we replace the GDP by the industrial production, table (2) shows that the different criteria selected 4 states and 1 or 2 lags. To be on the safe side we estimated the model with two lags, a MS-VAR(4,2). Table(4) shows the estimated mean for each variable in each one of the states and figure (6) shows the probability of the economy to be in each state.

Table(4) shows clearly that state 1 characterizes periods of deep recessions, as state 2 characterizes periods of mild recessions in which, despite the decrease in the industrial production and in the GFCF, the level of consumption growth at a slow rate. States 3

Table 4: Estimated mean over each Economy State

State	Ind. Prod.	Consumption	GFCF
		MS-VAR(4,2)	
1	-3.73910	-1.63026	-2.37701
2	-0.12847	0.18160	-1.16725
3	1.47077	0.90154	2.06225
4	2.86906	4.22257	0.15875

Figure 6: MS-VAR(4,2) States Probabilities



and 4 are periods of expansion, but these are different. While in state 3 the expansions can be seen as investment led, in state 4 they are clearly consumption led.

Figure (6) shows that states 2 (recession) and 3 (expansion) dominate the last 4 decades of the Portuguese economy. State 1 marked some quarters of the overall recession periods when the recession was more acute. State 4 (consumption led growth) was dominant until 1973, just before the democratic revolution in Portugal, marking the end of the pre-democratic expansionary period.

Taking together the two recession states , (1 and 2) the economy when characterized by the industrial production, consumption and GFCF was in recession in the following periods: 1974q2-1976q1; 1983q1-1985q3; 1991q1-1994q1; 2000q2-sample end. The main differences of this dating to the one in which the GDP is used instead of the industrial production are that previous to 2000 the recession periods were longer and the different recessions after 2000 are all merged in a long recession period, characterizing what many observers called the "lost decade"[7].

4.4 Results comparison, identification of recessions and annual correspondence

This subsection will compare and combine the results of the three used methodologies to propose a expansion, peak, recession, trough chronology. Afterwards, and because the salary data is only available at annual frequency we will make a correspondence between the quarter data and an annual chronology of the recession periods.

Table (5) summarizes the results of the previous methodologies, considering that a peak/trough is the last period of the previous expansion/recession.

Table 5: Recession Periods

Methodology					
Univariate				Multivariate	
BBQ		Markov-Chain		MS-VAR(5,1)	MS-VAR(4,2)
GDP	IP, C and GFCF	GDP	IP, C and GFCF	GDP, C and GFCF	IP, C and GFCF
74q1-75q2	74q1-75q2	74q2-75q3	74q2-75q4	74q2-75q2	74q2-76q1
80q1-80q3	80q1-80q3				
83q1-84q2	82q4-85q3	83q2-84q3	83q1-85q1	83q1-84q2	83q1-85q3
92q1-93q1	92q1-93q1	92q3-93q3	92q4-94q1	93q1-93q2	91q1-94q1
	94q4-96q4				00q2-...
	00q1-00q2				
02q2-03q2	01q4-05q3	02q2-03q3	02q1-05q3	02q2-03q1	
07q4-09q1	07q4-...	08q2-.09q3	08q2-.	08q2-09q2	
10q3-...		11q4-...		11q1-..	

From this table until 2000 the marking of the recessions is straightforward if we consider that when half or more of the estimation methodologies marked a recession. Therefore, the first recession went from 1974q2 until 1975q2 (even if some aggregates only recovered after 1975q2); the second recession took place from 1983q1 until 1985q1. The third recession went from 1992q1 until 1993q3, but this dating is heavily influenced by the BBQ methodology. Without it the recession would only start in the end of 1992 . Furthermore, as just the BBQ marked a recession in 1980 and in 1994q4-1996q4 we consider that there is no strong evidence to consider these two periods as recessions.

After 2000, the picture is messy. If we take the same approach as before, it can be considered that the recessions were from 2002q1 until 2005q3 and from 2007q4 until the sample end. However this result is due to the estimations that have used industrial production and not GDP. Therefore, this recession periods would dismiss the evolution of GDP and classify as recessions periods when the GDP was growing. If, on the

[7] See, between others, Lopes(2011) or Soukiazis and Antunes(2011)

contrary, we just use the GDP the first recession is shorter: 2002q2 until 2003q2 and the second is split into two: from 2008q2-2009q2 and from 2011q1 until the sample end. In fact if we look to the MC univariate methodologies we see that the consumption was almost never in recession until the very end of the period, so the short bursts of GDP growth that we observe in 2004-2005 and 2010 were mainly demand driven and not followed neither by industrial production nor investment. Even so, it would be difficult to consider these years as recession periods so we will classify them as demand driven expansions to distinguish them from the pure expansions when, also, GFCF and/or Industrial Production are growing. Table (6) summarizes our classification.

Table 6: Cycle chronology 72q1-12q3

Recessions	Expansions	Demand Expansion	Trough	Peak
...	72q1-74q1	...	...	74q1
74q2-75q2	...	...	75q2	...
...	75q3-82q4	...	...	82q4
83q1-85q1	...	...	85q1	...
...	85q2-91q4	...	...	91q4
92q1-93q3	...	...	93q3	...
...	93q4-02q1	...	...	02q1
02q2-03q3	...	...	03q3	...
...	...	03q4-05q3	...	...
...	05q4-07q3	...	...	07q3(IP)
...	...	07q4-08q1	...	08q1(GDP)
08q2-09q2	...	...	09q2(GDP)	...
...	...	09q3-10q4	...	10q4(GDP)
11q1-...	...	...	...	...

Finally, as the salary data is annual, we need to make a correspondence between these quarterly periods and the years. We will consider that a given year is classified as recession if 2 or more quarters were part of one of the recession periods identified above. Table (7) depicts the yearly chronology:

Table 7: Yearly chronology of recessions 1972-2012

Recessions	1974-75	1983-84	1992-93	2002-03	2008-09	2011-12

5 External economic conditions and labour market performance: the persistent effect of graduating in a recession

The analysis was conducted using data from the group of males, which facilitates the interpretation of the results by minimizing potential distortions such as discrimination. Due to the short span of available dataset, we will examine the issues only with cohorts entering the market before and during the recession year 1993. We hypothesize that new graduates who enter the labor force in a recession suffer a detrimental and persistent loss in comparison with the cohort who graduates in a boom. In particular, graduating in a slump might affect wage growth through two potential channels. First, since prevailing evidences suggest that job movers have higher returns to experiences than job stayers, recession may prevent job mobility so that slump cohort cannot recover from this setback by switching to better firms. Second, adverse economic shock may persist longer in wage path resulting in eventually less earnings than the boom group after a certain years being employed. We will investigate the issue in both static and dynamic analysis.

5.1 Static models

In this part, we will address the effects of recession on job transitions by exploiting the count and duration models. While the count data model is employed to explore whether entering in 1993 recession impinges the number of job switches, the duration model is used to compare the odds of work movements between two cohorts by accounting their employment spells. Three regression models for job counts are fitted into the data: the Poisson model, negative binomial model, and zero-inflated Poisson model. The conditional mean function can be written as:

$$E(y_i|X_i, u_i) = f(X_i\beta + u_i) \tag{1}$$

where yi is the number of job changes for individual i, Xi is the set of indicator of entry year, dummy of high education, the ratio of monthly wage over minimum wage, occupation, and size of associated firms. Details on these variables are summarized in the Table 1 below. The results are given in Table 2.

In all specifications, the coefficients on ENTERYEAR are significantly negative, which suggests that people who enter in 1993 are likely to switch jobs less than the 1992 cohort. When comparing the Poisson and negative binomial models, the test statistics fail to reject the hypothesis of equi-dispersion. Additionally, the descriptive data reveals little evidence on the existence of over-dispersion between the mean and

Table 8: Descriptive statistics

Variable	Mean	Std. Dev.	Min.	Max.
Number of jumps	0.50	0.722	0	5
Enter year (= 1993)	0.46		0	1
Educ (= high)	0.19		0	1
Age	21.8		18	30
LRatio	0.67	0.51	0.00	2.23
Size	4.63	2.04	0	10.29

Notes: For the descriptive statistics we use the sample with valid information on all relevant variables.

Table 9: Count data model for job changes

Variable	poison	negbin	zip
enteryear = 1993	-0.107	-0.108	-0.108
	(0.064)	(0.066)	(0.066)
educ=higher	0.007	0.005	0.053
	(0.126)	(0.128)	(0.112)
size	0.070	0.070	-.041
	(0.028)	(0.029)	(.205)
lratio	0.309	0.304	-.042
	(0.214)	(0.219)	(2.726)
size*lratio	-0.032	-0.031	-0.547
	(0.042)	(0.042)	(0.734)
N	1966	1966	1966
ll	-1819.82	-1819.13	-1822.73
Vuong			1.654

Notes: standard errors in parentheses.

variance of the number of jumps.Therefore, we would reject the negative binomial model in favour of Poisson model. On the other hand, the data reveals a preponderance of zeros (62% of the dataset).Thus Lambert's zero inflated model fits the data better in this case. In fact, Vuong diagnostic statistic confirms that the zero inflated model performs better than the standard Poisson regression; the value is 1.654.

Several duration models are employed to model labor market movements. In the initial step, we model the hazard rate of leaving employment which takes into account only time between entry and first job move (if any), and later we examine the hazard ratio of moving from jobs to jobs in the repeated event model accounting for the fact that individuals switch jobs more than once. In any cases, the hazard function is generally expressed as:

$$\lambda(t) = f(X_i\beta) \tag{2}$$

where X_i is the set of regressors including indicator of year entry, indicator of having high education, ratio of monthly wages over minimum wages, and size and location of the associated firms.

Concerning the hazard of leaving the first job, we conduct regressions on the 2 duration models: Weibull and Cox proportional model. Results are given in the following Tables 3 and 4.

Table 10: Estimated Hazard Ratio: Time until first move

	Weibull1	Weibull2	Cox1	Cox2
1993.enteryear	0.912*	0.958	0.917**	0.919*
	(0.035)	(0.042)	(0.029)	(0.033)
1.edu		1.564***		0.990
		(0.122)		(0.072)
size		0.997		0.975***
		(0.011)		(0.007)
lnratio		0.285***		0.893**
		(0.020)		(0.034)
Observations	2,785	2,262	2,785	2,262
Chi2	5.603	665.3	7.529	1499
LogLikelihood	-3680	-2620	-19528	-15214

Notes: standard errors in parentheses.; significance levels, *** p<0.001, ** p<0.01, * p<0.05.

Table 11: Estimated Coefficients: Time until first move

	Weibull1	Weibull2	Cox1	Cox2
1993.enteryear	-0.092*	-0.043	-0.086**	-0.085*
	(0.039)	(0.043)	(0.031)	(0.036)
1.edu		0.447***		-0.010
		(0.078)		(0.072)
size		-0.003		-0.025***
		(0.011)		(0.007)
lnratio		-1.256***		-0.114**
		(0.071)		(0.038)
Observations	2,785	2,262	2,785	2,262
Chi2	5.603	665.3	7.529	1499
LogLikelihood	-3680	-2620	-19528	-15214

Notes: standard errors in parentheses.; significance levels, *** p<0.001, ** p<0.01, * p<0.05.

The estimated parameters of ENTERYEAR in Weibull and Cox proportional models are quite similar. The significantly negative coefficient associated with ENTERYEAR suggests that 1993 graduates stay in their first job longer than the 1992 ones. In particular, entering in 1993 recession is likely to decrease 9% of likelihood of leaving the first job. Since choosing a particular distribution of baseline hazard, i.e. the individual heterogeneity, as the Weibull model could lead to biased hazard estimators while distribution for the Cox proportional model can get rid of this problem, the Cox proportional regression is more appropriate than the Weibull one. The estimated survival function and hazard function in the case of Cox model are reported in Figures 1 and 2.

We exploit the repeated event models to account for the possibility of recurrence events. It's reasonable to believe that for each individual the occurrences of moving

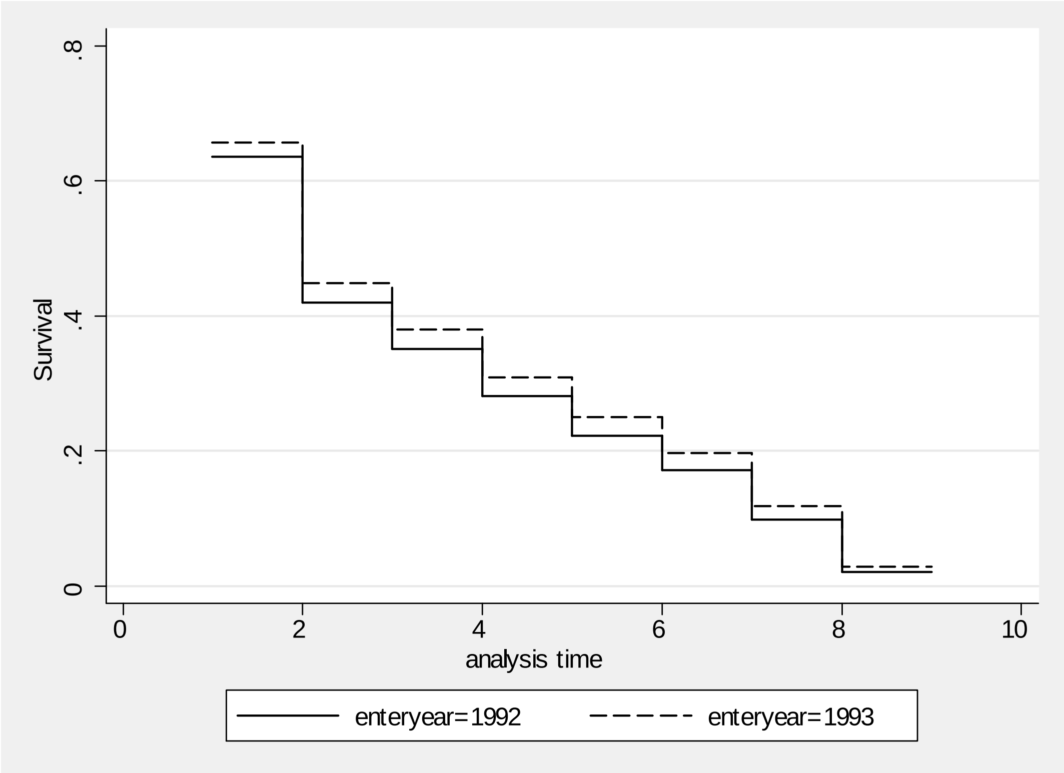


Figure 7: Survival: Cox proportional hazards regression

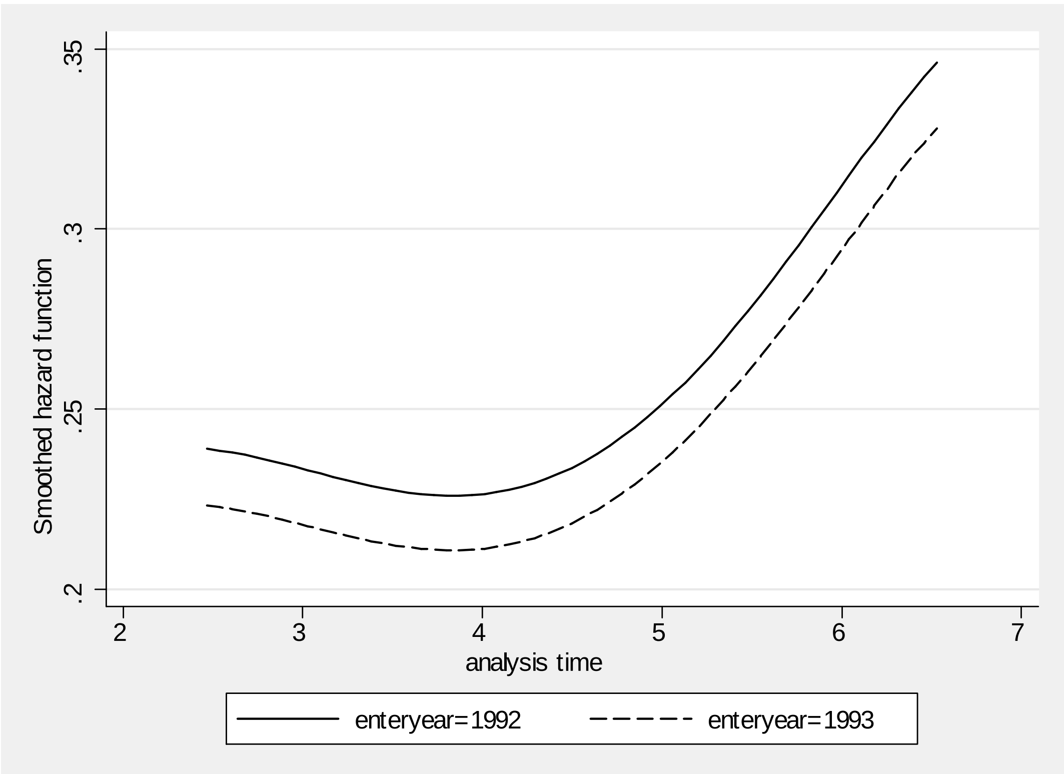


Figure 8: Hazard: Cox proportional hazards regression

(staying) are dependent in which either the second and subsequent events are influenced by the first event or there is a presence of unobserved heterogeneity across agents, for example, individual's ability. The differentiated ability over individuals might enable some particular types of agents to move from jobs to jobs more quickly than the others as several researches report that low educates have fewer opportunities to change jobs than the high educates while some others claim the reverse correlation. On the other hand, the existence of event dependence may come from the evidences that tenure reduces the hazard of changing jobs, i.e. the more likely one decides to stay in a firm today, the less likely he/she will move out of that firm tomorrow, or an increasing job changes leads to higher job quality through better matching efficiency. Hence, ignorance of event dependence in a multiple event model results in inconsistent estimates. In this paper, we therefore propose the Weibull frailty model, the standard Cox proportional model and the conditional risk set Cox model first presented by Prentice, Williams, and Perterson (1981) (hereafter PWP) to address these issues. The hazard function can be written as:

1. $\lambda_i(t) = \lambda_0(t)\gamma_i exp(X_{it}\beta)$ in the case of Weibull frailty model where γ_i is the frailty term assumed to follow a gamma distribution in Weibull model
2. $\lambda_i(t) = \lambda_{0i}(t)exp(X_{it}\beta)$ in the case of standard Cox model
3. $\lambda_{ik}(t) = \lambda_{0k}(t - tk - 1)exp(X_{it}\beta)$ where k denotes the event number and λ_{0k} is the baseline hazard for event k, $t - tk - 1$ incorporates the gap time structure so that the hazard gives the risk for event k since $(k-1)th$ event.

In these 3 models, X is the set of covariates including indicator of entry year, years of education, age, ratio of monthly wages over minimum wages, and size and location of associated firms. Table 4 and 5 give the regression results.

Table 12: Estimated Coefficient: Repeated event models

	Frailty1	Frailty2	Frailty3	Cox1	Cox2	Cox3	PWP1	PWP2	PWP3
1993.enteryear	-0.057	-0.030	-0.017	-0.094***	-0.088***	-0.041	-0.023	-0.009	-0.012
	(0.056)	(0.056)	(0.039)	(0.005)	(0.024)	(0.031)	(0.083)	(0.079)	(0.080)
1.edu		0.927***	1.395***		0.223***	1.077***		0.026	0.876***
		(0.083)	(0.083)		(0.046)	(0.076)		(0.237)	(0.304)
lratio		-1.841***	0.082		-0.641***	0.168***		-1.579***	-0.849***
		(0.069)	(0.063)		(0.038)	(0.046)		(0.237)	(0.200)
size		0.002	0.014		0.006	0.006		-0.755***	0.074
		(0.013)	(0.010)		(0.006)	(0.007)		(0.242)	(0.367)
age			-0.533***			-0.558***			-0.364***
			(0.012)			(0.017)			(0.042)
Observations	4,176	4,176	3,489	4,176	4,176	3,489	2,238	2,238	2,238
theta	0.786	0.866	0.110						
Chi2	1.041	905.8	3033	421.4	381.0	2000	0.0806	52.44	137.3
LogLikelihood	-4993	-4540	-2694	-29264	-29137	-22395	-523.7	-509.1	-482.7

Notes: Robust standard errors in parentheses; significance levels, *** $p<0.01$, ** $p<0.05$, * $p<0.10$

Table 13: Estimated Hazard ratio: Repeated event models

	Frailty1	Frailty2	Frailty3	Cox1	Cox2	Cox3	PWP1	PWP2	PWP3
1993.enteryear	0.945	0.971	0.983	0.910***	0.916***	0.960	0.977	0.991	0.988
	(0.053)	(0.054)	(0.039)	(0.004)	(0.022)	(0.030)	(0.081)	(0.078)	(0.079)
edu		2.527***	4.036***		1.250***	2.935***		1.026	2.402***
		(0.209)	(0.336)		(0.058)	(0.222)		(0.243)	(0.729)
lratio		0.159***	1.085		0.527***	1.183***		0.206***	0.428***
		(0.011)	(0.068)		(0.020)	(0.055)		(0.049)	(0.085)
size		1.002	1.014		1.006	1.006		0.470***	1.077
		(0.013)	(0.010)		(0.006)	(0.007)		(0.114)	(0.396)
age			0.587***			0.572***			0.695***
			(0.007)			(0.009)			(0.029)
Observations	4,176	4,176	3,489	3,489	4,176	3,489	2,238	2,238	2,238
theta	0.786	0.866	0.110						
Chi2	1.041	905.8	3033	421.4	381.0	2000	0.0806	52.44	137.3
LogLikelihood	-4993	-4540	-2694	-29264	-29137	-22395	-523.7	-509.1	-482.7

Notes: Robust standard errors in parentheses; significance levels, *** p<0.01, ** p<0.05, * p<0.10

The log-likelihood ratio tests and Akeike information criterion report that the full regressions of Weibull frailty model, standard Cox model and conditional risk set Cox model are preferable to their constrained models. The estimated thetas in frailty models are significantly different from 0, suggesting that there is presence of unobserved heterogeneity. While the frailty model corrects for heterogeneity by positing a random effect factor into the survival function, the Cox proportional model addresses the group- or individual effects by adjusting the variance-covariance matrix, similarly the way we deal with heterogeneity in OLS models. However, these models do not take into account the possibility of event duration in which they assume that the occurrences of events for each unit are independent. As we believe that the decisions of sequential job switches correlate with the first switch, the problem of event dependence raises our concern. In fact, employing the approach of Prentice, Williams, and Peterson(1981) with allowing distinct baseline hazard ratios for different specific types of agents through setting up stratification , we can account for the issues of both event correlation and individual heterogeneity.

The results of PWP models, again, reveal that entering in 1993 recession corresponds to around 12% decrease in the hazard of job moving. Besides, significantly negative coefficient associated with AGE implies that older people have less incentive to change jobs than the younger. We also find that people who earn higher wages will be less likely to leave the job as the coefficients of RATIO are negative. The results also reveal that the relatively bigger enterprises one works, the less odd he/she will have job transition; however, the coefficients associated with SIZE are not consistently negative in all specifications. Interestingly, the positive coefficient associated with EDU suggests that higher education people are likely to retain in their jobs shorter than the lower educates. Particularly, according to the PWP model 3 the likelihood of leaving jobs is about 82% lower in the case of attaining low education.

5.2 Dynamic models

We can illustrate the idea using the simple first order equation

$$w_t^i = \beta w_{t-1}^i + \alpha + \varepsilon_t^i \tag{3}$$

where $\{w_t\}_{t=0}^{\infty}$ is the sequence of real wages of each individual of type i overtime; $\{\varepsilon_t\}_{t=1}^{\infty}$ is a sequence of random shocks that have an expected value of zero. We assume that $|\beta| < 1$.

The general solution of the equation (1????) has the form:

$$w_t^i = \frac{1-\beta^t}{1-\beta}\alpha + \beta^t w_0^i + \sum_{i=0}^{t-1} \beta^i \varepsilon_{t-i} \tag{4}$$

According to the form of solution (2), the more persistent the autoregressive factor is, i.e. higher β, the greater and longer the effects of shocks on the sequence of wages are. Hence, by comparing the coefficient associated with the lag wage we can detect to what extend the aggregate shock can affect the wage path.

5.2.1 Persistent effects of graduating in a recession

Target the young male individuals who enter the market in either 1992 or 1993

The question we examine here is whether the average wage level of the male cohort who enters the workforce during the recession year 1993 evolves more persistently than the average wage of the ones entering the market in a boom year 1992. In particular, we are going to examine the four following models but taking into account the results from the full model, i.e. the 4th one.

$$lnwage_{it} = \beta_0 + \beta_1 lnwage_{it-1} + \eta_i + Y_t + \varepsilon_{it} \tag{5}$$

$$lnwage_{it} = \beta_0 + \beta_1 lnwage_{it-1} + \beta_2 lnsale_{it} + \eta_i + Y_t + \varepsilon_{it} \tag{6}$$

$$lnwage_{it} = \beta_0 + \beta_1 lnwage_{it-1} + \beta_2 lnsale_{it} + \beta_3 occup_{it} + \eta_i + Y_t + \varepsilon_{it} \tag{7}$$

$$lnwage_{it} = \beta_0 + \beta_1 lnwage_{it-1} + \beta_2 lnsale_{it} + \beta_3 occup_{it} + \beta_4 ind_{it} + \eta_i + Y_t + \varepsilon_{it} \tag{8}$$

Where *lnwage* is either the log real hourly wage or log real hourly base wage of employee i at year t; *lnsale* is the log profit of the firm that the employee i is working each year. Since individuals are observed to switch their jobs and their industries over time, we include occup and ind the time-variant dummy variables for current occupation and industry, respectively, into the model as potential explanatory variables. Unobserved error components include a calendar year effect Y_t , individual effect η_i and the remaining error ε_{it} .

In the absence of serial correlation, the GMM estimated will be efficient and consistent. Since both past wages and firm's profit are endogenous variables, we introduce a set of lagged values of lnwage and lnsale as instrument variables to correct for endogeneity. Moreover, individual's occupation and industry are assumed to be exogenous.

The key parameter of interest is β_1 capturing the endogenous persistence of wage evolution. Both estimates of the system GMM model and FE and OLS estimates are reported in the following Table 6 and Table7 :

The reported FE and OLS estimates identify the range in which the parameter estimates of the dynamic model should lie if correctly specified. The estimates β_i in the 1993 graduate group is significantly higher than that of the1992 graduate group meaning that the wage path of the cohort entering the market in the recession are likely to be more persistent than that of the cohort graduating in the preceding year. Moreover, if the 1993 graduates are on impact more exposed to recession shocks than the 1992 ones, this effect will last longer with greater magnitude in the case of the 1993 cohort.

To gain some insight into the effects of graduating in the recession on wages, we consider the movement of the ratio between each cohort's wages over minimum wage overtime. By using minimum wage as the frame of reference, the ratio will tell us how fast these graduates are able to move up the job ladder. We also conduct the following 4 regression to uncover the query in question:

$$lnratio_t = \rho_0 + \rho_1 lnratio_{t-1} + \eta_i + Y_t + \varepsilon_{it} \tag{9}$$

$$lnratio_t = \rho_0 + \rho_1 lnratio_{t-1} + \rho_2 lnsaleit + \eta_i + Y_t + \varepsilon_{it} \tag{10}$$

$$lnratio_t = \rho_0 + \rho_1 lnratio_{t-1} + \rho_2 lnsaleit + \rho_3 occupit + \eta_i + Y_t + \varepsilon_{it} \tag{11}$$

$$lnratio_t = \rho_0 + \rho_1 lnratio_{t-1} + \rho_2 lnsaleit + \rho_3 occupit + \rho_4 ind_{it} + \eta_i + Y_t + \varepsilon_{it} \tag{12}$$

Where lnratio is the ratio between nominal monthly wages and minimum wages each year. The other variables remain as discussed above.

The key parameter estimate of interest is ρ_1 explaining how persistently the wages rise over the minimum wages. Both estimates of the system GMM model and FE and OLS estimates are reported in the following Table 8 and 9:

The estimate ρ_1 in the 1993 group is higher than the one of the 1992 group. It means that graduates in a recession are likely to face a slower wage growth in comparison with the other group. In other words, they get "stuck" in the same-paid jobs longer and experience slower promotion than the graduates in the other group. We next compare the significance degree of the disparities between each pair of parameter estimates in concern by considering these two models:

$$lnwage_{it} = \beta_0 + \beta_1 lnwage_{it-1} + \beta_2 lnsale_{it} + \gamma inter_i + \beta_3 occup_{it} + \beta_4 ind_{it} + \eta_i + Y_t + \varepsilon_{it} \quad (13)$$

$$lnratio_t = \rho_0 + \rho_1 lnratio_{t-1} + \rho_2 lnsale_{it} + \lambda inter_i + \rho_3 occup_{it} + \rho_4 ind_{it} + \eta_i + Y_t + \varepsilon_{it} \quad (14)$$

Where inter in the model 3 and 4 are created by interacting of the log real wages and log ratio, respectively, with an indicator of recession, wherein the indicator is given value 1 if the individual enters labor force in 1993 and 0 if otherwise. The other variables remain unchanged. The estimates in place γ and λ are designed to assess the magnitude and the degree of significance of each pair of the estimates associated with lnwage and lnratio from two different cohorts. The results are reported in the Table 10 and 11:

The interaction in table 9 reports that the coefficients associated with the lag wage in the group 1993 are significantly around 6% higher than ones with the lag wage of group 1992 while this gap, reported in table 10, is significantly around 12% when using the log ratio of wage.

Target the young male individuals who enter the labor force in 92 or 93 and who are either low educated or high educated

To uncover the question which types of workers are most affected when graduating in the 1993 recession, we employ the analogous analysis to the groups differentiating in level of education. We separate the young male group into the subgroup of high education who acquire more than 12 years of education before entering the labor force and the subgroup of low education who attain 12 years of education at most. The table 11 and 12 below report the estimates of interest of SYS-GMM models in the low and high education subgroup, respectively.

Again, the estimated coefficients associated with the low (high) educated graduates with 1993 entry year are higher than those of the group who enters in 1992. It implies that the individuals entering labor force in 1992 belonging to either the high or low education group are likely to recover more quickly when shock is taken and to suffer the less severe setback than the other group if we are concerned that the magnitude of the 1993 shock is substantially greater than of the 1992 shock.

We now re-examine the above mentioned evidence by testing the models with interaction terms. First, we simply employ the regression on models with an interaction variable, inter, for each sub group, wherein log real wage is interacted with 1993 entry dummy. The interaction variable is subject to examine the discrepancies of monthly wage growth between the two cohorts. The significantly positive values of the interaction variable in the case of low education group inter confirm our hypothesis; however, the estimates of the interaction variable in the case of high education are not. Results are reported in the table 13 and 14 as follows:

We then apply the difference-in differences method to see how the advantage of being highly educated was impacted due to the recession. Namely, we generate 3 dummy variables inter1, inter2, inter3 representing for the "graduating in a recession' treatment effect, the different group effect between high and low education and the difference-in differences effect of graduating in a recession on higher education advantages. In details, inter1 is the interaction between log real hourly wage and 1993 entry dummy, inter2 is resulted from log real hourly wage interacted with high education dummy and inter3 is the triple interaction of log real hourly wage, 1993 entry dummy and high education dummy. Regression results are displayed in table 15 below:

The coefficient of inter1 is significantly positive, as expected. It confirms that the persistence of wage of the 1993 entry group is higher than the persistence of wage of the other group due to the graduation year. The coefficient of inter3 is significantly negative implying that the wage disparity between the high and low educates wages is higher when there is an occurrence of recession, meaning that the low educates are impinged more than the high educates. Interestingly, we obtained consistently significantly positive values of inter2 which implies that wage path of the high educated is, instead, more persistent than the path of low educated group. Put differently, the effects of shocks persist longer on wage path in the case of high education group; yet, inter3 told us that the magnitude of these effects is smaller in the case of high educates. This result is consistent with the findings in our static model. Since high educates are more likely to switch jobs, they can recover better than the low educates by job mobility. Hence, the magnitude of shock effects on high educates' wage path is smaller.

6 Conclusion

The recent global economic crisis recalled the need to better understand the impact of recessions on the labour market and on the social dimensions of our societies. The fact that countries throughout the world have been affected in different years and for different periods suggested the need to firstly identify the periods of booms and busts. This exercise has allowed to identify the Portuguese business cycle providing the relevant period for the study of the impact of recessions in the Portuguese labour market.

Motivated by a rich strand in the economic literature and the persistently high inequality indicators found in Portugal, the relation between the business cycle and income inequality was investigated. The inconsistency of preliminary results suggested that the impact of the business cycle on income inequality could be hidden in the impact of the business cycle on other labour market dynamics and therefore a deeper analysis was required. Again, following the literature, the impact of graduating in a recession was evaluated in terms of initial wages and their persistence throughout more than a decade. The results suggest that there are important negative consequences of graduating in a recession and that these consequences differ across different educational levels.

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Appendixes

A A short description of the Markov Chain univariate methodology[8]

In the first part of the appendix we will describe the underlying Markov Chain (MC) and the second part deals with the scoring of its transition probabilities.

A.1 The underlying Markov Chain

At any given time t, the economy is in either of two mutually exclusive states: expansion, denoted by E_t or recession, R_t in which the last period of an expansion is a peak (P_t) and the last period of a recession is a trough (T_t). So:

$$E_t = \begin{cases} EC_t & \text{(Expansion continuation)} \\ P_t & \text{(Peak)} \end{cases}$$
$$R_t = \begin{cases} RC_t & \text{(Recession continuation)} \\ T_t & \text{(Trough)} \end{cases}$$

From this definition from an EC_t we can continue the expansion (EC_{t+1}) or make a transition to a peak (P_{t+1}). But from a peak we cannot go back to a expansion, the only admissible state at $t+1$ will be a recession continuation ($P_t \rightarrow RC_{t+1}$ with probability one)[9]. The same happens if the economy at t is at the state RC_t. So denoting by $p_{EP} = P(P_{t+1}|EC_t)$ the probability of making a transition to a peak and $p_{EE} = P(EC_{t+1}|EC_t) = 1 - p_{EP}$ the probability of continuing the expansion and analogously to the recession: $p_{RT} = P(T_{t+1}|RC_t)$ and $p_{RR} = P(RC_{t+1}|RC_t)$ we define a first-order MC with four states, denoted S_t, with transition matrix:

	EC_{t+1}	P_{t+1}	RC_{t+1}	T_{t+1}
EC_t	$p_{EE} = 1 - p_{EP}$	p_{EP}	0	0
P_t	0	0	1	0
RC_t	0	0	$p_{RR} = 1 - p_{RT}$	p_{RT}
T_t	1	0	0	0

[8] This appendix presents a brief description of the methodology. For more details refer to Artis et al.(2004) where it was first proposed.

[9] Note that we will impose that each state (Expansion or Recession) and one complete cycle will have a minimum duration constraint, so we cannot go from a peak to a trough immediately.

The dating rules impose a minimum duration phase of two quarters[10] and a minimum duration of a cycle (a complete expansion and a complete recession phase) which amounts to five quarters. Note that these duration are a direct transposition of Bry and Boschan (monthly) rules (of respective 6 and 15 months) applied to quarterly data. Note that the fact that a complete cycle is composed by a complete expansion and a complete recession phase excludes the following pattern $\{P_{t-4}, RC_{t-3}, T_{t-2}, EC_{t-1}, P_t\}$ as the peak at time $t-4$ is part of the previous expansionary phase, the complete cycle is comprised between $t-3$ and t, for a total of four periods violating our 5 quarters minimum duration rule.

The tie on the full cycle duration yields a fifth-order MC that can be transposed to a first-order one by combining the elements of the original chain, S_t. The states of the new MC are defined by:

$$S_t^* = \{S_{t-4}, S_{t-3}, S_{t-2}, S_{t-1}, S_t\}$$

The ties described before conduce that the number of possible states (S_t^*) in the derived MC are 24, and the number of states to which each one of these states can transit ($S_t^* \to S_{t+1}^*$) varies between 1 and 2, and whenever the possible transitions are 2 the probabilities are given by one of the following pairs $\{1-p_{EP}, p_{EP}\}$ or $\{1-p_{RT}, p_{RT}\}$. So the two parameters p_{EP} and p_{RT} uniquely specify the MC.

A.2 Scoring the transition matrices

Once we know p_{EP} and p_{RT} the MC is fully specified. These two parameters can be estimated by maximum-likelihood, if it is assumed that the series are a realization of a stochastic process that is dependent of the state of the economy as is represented by the chain. However, Artis et. all(2004) when proposed this methodology adopted a different strategy and use a non-parametric approach[11] suggested by Harding and Pagan(2002), according to which an expansion-termination sequence (ETS) and a recession termination sequence (RTS), are defined, respectively by:

$$\begin{aligned} ETS_t &= \{(\Delta y_{t+1} < 0) \cap (\Delta_2 y_{t+2} < 0)\} \\ RTS_t &= \{(\Delta y_{t+1} > 0) \cap (\Delta_2 y_{t+2} > 0)\} \end{aligned}$$

Here Δ is the backward difference operator $\Delta y_{t+1} = y_{t+1} - y_t$. The joint distribution of the sequences ETS_t and RTS_t for $t = 1, 2, ..., T$ depends on the stochastic process generating the series and is usually analytically intractable, due to the presence of

[10]Note that this first duration rule is automatically enforced by the four states characterization as a peak cannot follow a trough and vice-versa.

[11]For a comparison with the parametric approach see Harding and Pagan(2003).

autocorrelation and the fact that the termination processes are not mutually exclusive. As regards the latter, denoting by $\overline{ETS}_t$ the complementary event of ETS_t, $\overline{RTS}_t$ that of RTS_t, and defining $P_t^{ETS} = P(ETS_t), P_t^{RTS} = P(RTS_t)$, the joint probability distribution of the possible occurrences at time t is:

	ETS_t	$\overline{ETS}_t$	$Marginal$
RTS_t	0	P_t^{RTS}	P_t^{RTS}
$\overline{RTS}_t$	P_t^{ETS}	$1 - P_t^{ETS} - P_t^{RTS}$	$1 - P_t^{RTS}$
$Marginal$	P_t^{ETS}	$1 - P_t^{ETS}$	1

hence it can be seen that ETS_t and RTS_t cannot both be true at the same time.

Serial correlation complicates the computation of P_t^{ETS} and of P_t^{RTS}, as the terminating sequences are not independent of their past; furthermore, it must be stressed that the BBQ rule induces autocorrelation itself. Therefore, stochastic simulation is the only way to go about the characterization of the business cycle for a particular stochastic process.

Let us return to the non-parametric scoring of the transition probabilities cording to the available time series. If at time t, the chain S_t^* is in any of the expansionary states for which a transition to a peak is possible and an expansion terminating sequence occurs at time $t+1$, i.e. ETS_{t+1} is true, then we move to a new state: S_{t+1}^* such that $S_{t+1} = P_{t+1}$. So we can divide the states of S_t^* into the following subsets: $\Phi_{EP,}$which includes the states that have a expansionary state at time t and are available for a transition peak; $\Phi_{P,}$which includes the states that have a peak state at time t; $\Phi_{RT,}$which includes the states that have a recession state at time t and are available for a transition trough; $\Phi_{R,}$which includes the states that have a trough state at time t. Define also $\Phi_{E,}$ and Φ_R the sets with expansionary and recession states at time t.

Algorithm:

1. If $\{S_t^* = s_{EP}, s_{EP} \in \Phi_{EP}\}$ and, ETS_{t+1} is true then $\left\{S_{t+1}^* = s_P, s_P \in \Phi_P\right\}$. Hence, the transition probability is computed as:

$$p_{EP} = P\left(\{S_t^* = s_{EP}, s_{EP} \in \Phi_{EP}\} \cap ETS_{t+1}\right) = I(ETS_{t+1}) \sum_{s_{EP} \in \Phi_{EP}} P\left(S_t^* = s_{EP}\right)$$

where $I(.)$ is the indicator function. Else, if ETS_{t+1} is false, expansion continues and $p_{EE} = 1 - p_{EP}$.

Else, if $\{S_t^* = s_{RP}, s_{RP} \in \Phi_{RP}\}$ and, RTS_{t+1} is true then $\left\{S_{t+1}^* = s_T, s_T \in \Phi_T\right\}$. Hence, the transition probability is computed as:

$$p_{RT} = P\left(\{S_t^* = s_{RT}, s_{RT} \in \Phi_{RT}\} \cap RTS_{t+1}\right) = I(RTS_{t+1}) \sum_{s_{RT} \in \Phi_{RT}} P\left(S_t^* = s_{RT}\right)$$

Else, if RTS_{t+1} is false, recession continues and $p_{RR} = 1 - p_{RT}$

Probabilistic dating based on a maintained stochastic process replaces the indicator functions by the probability of terminating sequences: .P_t^{ETS} and P_t^{RTS}.

Let $\mathcal{F}_t$ denote the collection of $I(ETS_j)$, $I(RTS_j)$ and $j = 1, 2, ..., t$ and $P(S_t^*|\mathcal{F}_t)$ denote the probability of being in any particular state at time t, conditional on this information set. Assuming that this probability is known, we can recursively compute the probability of the chain at subsequent times by using the following filter:

1 Given the availability of $P(S_t^*|\mathcal{F}_t)$ at time t, let us denote by π_t^* the $m \times 1$ vector containing them, with $m = 24$ in the quarterly case. Define the two $m \times 1$ selection vectors v_{EP}, with 1 corresponding to the elements of Φ_{EP} and 0 otherwise, and v_{RT}, with 1 corresponding to the elements of Φ_{RT} and 0 otherwise.

2 Compute the transition probabilities of the chain according to the two previous equations and insert them in the transition matrix of the chain, hereby denoted by Γ..

3 Compute the probabilities $P(S_{t+1}^*|\mathcal{F}_{t+1})$ belonging to the vector π_{t+1}^* as $\pi_t^* = \Gamma'\pi_t^*$

The algorithm is initialized by assigning values to π_1^*, where we can input any information we know about the first period or if we do not want or if the information is ambiguous we can use a uniform prior, which amounts to setting the elements of π_t^* equal to $1/m$.

The algorithm recursively produces $P(S_t^*|\mathcal{F}_t)$, for all $t = 1, \ldots, T$, and hence, marginalizing previous states, the probabilities of each elementary event.