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Smooth transition trends and labor force participation rates in the United States

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Abstract This paper employs smooth transition trend models to investigate the long-run time series behavior of quarterly US labor force participation rates. In particular, we examine whether long-run growth in labor force participation rates can be modeled by smooth transitions between states rather than as abrupt mean level changes or as a stochastic trend. Smooth transitions permit for non-instantaneous adjustment of individual workers to changes associated with economic events or general labor market conditions. We employ unit root testing procedures with alternatives characterized by stationary fluctuations around one or two smooth transitions in linear trend. We examine labor force participation rates by gender- and age-specific groups. The results indicate that all female and most male participation series are better characterized as stationary processes that undergo transitional deterministics.

Keywords Labor force participation rate · Non-linear trend · Deterministic smooth transition · Unit root

JEL classification C22 · J21

1 Introduction

Labor force participation rate in the United States has undergone salient changes in the postwar era, with female rates exhibiting the most prominent movements. The labor force participation rate of women has risen remarkably in the postwar period, while a modest decline has been observed in male participation rates. Understanding such

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changes in labor force participation rates has attracted a considerable amount of work in the field of labor economics.

Potential explanations and causes of female labor force participation include demographic and social factors, such as declined fertility and delayed marriage, as observed by a rising age at first marriage, higher divorce rates, and increased educational attainment, with the latter being considered as one of the most crucial factors that led into higher female participation rates. Killingsworth and Heckman (1986) provide evidence of a strong positive relationship between educational attainment and labor force participation rate, implying that higher levels of education are associated with increased probability of entrance in the labor market. In addition, changes in social attitudes encouraged the entrance of married women (possibly with young children) in the labor market, which further contributed to the increase in the aggregate female labor force participation rate.

Further explanations include increasing acceptance of women in the labor market and rising wage rates. In particular, there has been an increase in the demand of women's labor in some sectors of the economy (e.g., service sector), which further resulted in rising real wages for women. Some additional explanations include changes in household technology and the introduction of labor-saving devices which reduced substantially the amount of time spent on household (Greenwood et al. 2005) and changes in husband's income.

On the other hand, the modest decline in males' participation rates is mainly attributed to the steep decline in the participation of older men near or above 65 related to the introduction of Social Security and early retirement of older workers, and the reduction in the participation rate of prime-age males, i.e., those aged 25–54. This reduction has been attributed to the decline in demand of less-skilled workers—primarily due to technical change—which further resulted in decreasing wages of less-skilled men and a steep fall in less-educated men's employment (Juhn 1992; Murphy and Topel 1997). Moreover, except for early retirement, the declining demand of less-skilled workers appears to have affected participation of older workers as well (Peracchi and Welch 1994). Comprehensive reviews of these long-term trends in the US labor supply may be found in Michael (1985); Smith and Ward (1985); Killingsworth and Heckman (1986); Pencavel (1986); Jacobsen (1999); McEwen et al. (2005); Juhn and Potter (2006) and DiCecio et al. (2008).

Variations in the labor force participation rate are also related to business cycles, with recessions being often responsible for pauses or reductions in the participation rate. This observation is associated to the "discouraged worker" notion, which refers to workers' tendency to give up searching for work during a recession and drop out of the labor force. Some empirical studies that have investigated the existence of discouraged worker effects using either macro time series data or microdata are Benati (2001); Darby et al. (2001) and Dagsvik et al. (2012).

In recent years, the participation behavior has regained the attention of labor economists following Murphy and Topel (1997) who, using data on American males since the mid-1960s, find "discouraged" demand-driven worker effects, mostly among the least skilled men that led them to conclude the debilitation of the unemployment rate as an informative indicator of joblessness and the state of the labor market. In particular, they find that declining labor market opportunities resulting from long-term changes in labor demands have led less skilled men to withdraw the labor market rather than entering the unemployed. As a result, since individuals (i.e., workers) may drop out of the labor force for market-driven reasons and unemployment data lack the non-employed or out of the labor force workers, it has become evident that the unemployment rate alone, without examining participation behavior, cannot further constitute a reliable indicator of labor market conditions (see also Juhn and Potter 2006).

Several recent empirical studies have concentrated on the participation behavior by examining the time series properties of labor force participation rates.[1] Moreover, the characterization of trends in the labor force participation is important not only from a time series perspective but also for the way labor force participation is introduced to more advanced models, for example VAR models that examine time series causality issues.

Gustavsson and Österholm (2006) are the first to apply standard univariate unit root tests as well as panel unit root tests to examine the (non)stationary nature of participation rates. In particular, they employ the Augmented Dickey-Fuller test of Said and Dickey (1984), the Augmented Dickey-Fuller test with GLS detrending of Elliott et al. (1996), the Kapetanios et al. (2003) KSS test, and the Kwiatkowski et al. (1992) KPSS test along with the panel unit root tests of Im et al. (2003) and Johansen (1988) likelihood ratio tests. They use monthly aggregate data for Australia, Canada, and the US with samples that begin in February 1978, January 1976 and January 1951 respectively and cover the period up to November 2004. They conclude in favor of unit root non-stationarity. Gustavsson and Österholm (2012) apply univariate unit root tests to disaggregated participation rates of subpopulations of the US labor force (disaggregation by gender, race, age) and find no evidence of stationarity. In both studies, more complicated trend alternatives that allow for structural breaks or non-linear trends have not been investigated.

Madsen et al. (2008) take into consideration evidence of asymmetric labor force participation responses to different labor market conditions and apply the Caner and Hansen (2001) unit root tests in the presence of a non-linear stochastic threshold. Non-linear behavior is rationalized by observed differences in individuals' labor supply responses under different economic conditions, e.g., when employment opportunities weaken from when they improve (see also Hotchkiss and Robertson 2012). They further use the Lagrange Multiplier (LM) unit root test with one and two breaks in the trend intercept and slope developed by Lee and Strazicich (2003, 2004) to examine whether participation rate series are trend reverting. They use annual data over a period of 130 years for G7 countries, with the results from both methods being mixed.

Landajo and Presno (2010) develop a test with the null of stationarity around generic smooth trend functions approximated by trigonometric series which can be applied to series that may display non-linear trend behavior, such as the labor force participation

[1] Most of these studies question on the informational value of unemployment rates or the presence of unemployment hysteresis (see the relevant study of Gustavsson and Österholm 2010), with evidence of non-stationary participation rates implying hysteresis in unemployment and uncertainty about the relation between long-term changes in unemployment and employment rates (see Madsen et al. 2008 for further discussion).

rate. They employ monthly data for the US labor force participation rate between January 1948 and August 2007 and find evidence of non-linear trend stationarity. Finally, Ozdemir et al. (2012) test for unit roots while allowing endogenously determined (instant) multiple structural breaks in labor force participation rates. They use the Robinson (1994) LM unit root test and the procedures of Gil-Alana (2008) and Bai and Perron (1998) for determining multiple structural breaks at unknown dates. They extend the monthly data of Gustavsson and Österholm (2006) to July 2008 and employ total and disaggregated participation rates by gender and conclude that structural trend breaks render the total, female, and male participation rate series stationary or at best mean reverting.

In general, changes in economic aggregates depend on changes in the behavior of a large number of agents. Besides, the reaction of individual agents to general economic events is unlikely to occur simultaneously. In contrast, some agents may be able (and want) to adjust instantaneously, while others will adjust in different time horizons, leading to smooth or gradual aggregate changes in economic time series. Where labor market is concerned, gradual adjustment of participation rates appears reasonable when aggregate behavior is considered. The reaction of individual agents (with respect to their participation decision) to changes in labor market conditions, government policies or general changes in the economic environment is not expected to be uniform either across agents of a particular group or across different demographic groups. For example, the growth in the participation rate of women began at different times and continued with different growth rates across individual age groups or groups with different levels of education. Similar patterns are observed for male participation rates.

Importantly, the kinds of factors that have influenced the movements in US labor force participation rates may be sufficient by themselves to rationalize the aspect of gradual or smooth adjustment in the participation series (see Van Zandweghe 2012). For example, the social or demographic factors or the improvement in labor-saving housing technologies have not instantly affected the (aggregate) labor force participation rate of women, rather they have induced a continuous rise in females' participation rate that lasted for a long period of time. Similarly, the reduction in the participation rate of less-skilled men due to skill-biased technical change or in the participation rate of older men because of changes in the Social Security did not occur rapidly but continued for several years as well.

In this paper, we aim to shed new light on the time series properties of labor force participation rates in the United States. Based on the notion of smooth aggregate changes in economic time series and particularly on the argument that the time path of structural changes in the aggregate labor force participation rate may be better captured by models whose deterministic component permits gradual rather than instantaneous adjustment between states, we employ smooth transition (STR) trend models to investigate the long-run evolution of participation. This approach presumes that long-run movements in participation rates may be more appropriately modeled as smooth transitions between states or regimes rather than as abrupt mean level changes or as a stochastic trend. Under the smooth transition specification, labor force participation rates are better characterized as stationary fluctuations around a deterministic nonlinear component, rather than as non-stationary unit root fluctuations that embody the permanence of stochastic shocks, while presenting no tendency to revert to a long-run mean level.

Our analysis employs the unit root tests proposed by Leybourne et al. (1998) and Harvey and Mills (2002). Leybourne et al. (1998) make use of the logistic smooth transition function and introduce tests that allow a gradual deterministic transition between regimes under the alternative. The transition midpoint and speed are determined endogenously, while the test procedure, under the alternative, can also provide information (a) on the range over which the transition takes place, (b) on the relative transition speed by recording the duration of the transition path and (c) on the direction of the transition. The unit root tests of Harvey and Mills (2002) extent the previous procedure such as the alternative admits two structural changes in the deterministic trend.

We examine labor force participation rates by gender- and age-specific groups. In particular, we employ both male and female age-specific participation rates to draw conclusions on the presence of smooth transitions across different labor force participation rate components. Both gender and age are important factors which induce heterogeneity across agents (workers), thus differences in the time series characteristics are not unexpected.

Our results indicate that the permission of a more versatile trend function when testing for unit roots provides substantial evidence against the hypothesis that the long-run evolution of participation rates is best described by a stochastic trend. We find that most labor force participation rates are better characterized as stationary processes that undergo transitional deterministics. Evidence from the estimated transition midpoints as well as the duration and direction of the smooth transitions is in full agreement with the movements of labor force participation rates in the United States.

The paper is organized as follows. In Sect. 2, we present the smooth transition (STR) trend models that serve as alternatives to unit root behavior in labor force participation. Section 3 offers a brief description of US labor force participation trends and thereinafter presents the estimation and testing methodology and discusses the empirical results. Finally, Sect. 4 concludes.

2 Smooth transition trend alternatives

The use of smooth transition analysis as an approach for representing deterministic structural change in time series regression has been well established in the work of Granger and Teräsvirta (1993) and Lin and Teräsvirta (1994). Bacon and Watts (1971) first suggested a model to represent a smooth rather than a discrete transition from one regime to another, in which the regressor was also the transition variable. From the structural change perspective, the transition variable is a function of time.

The basic idea of smooth transition models (STR) has been built upon the argument that the reaction of individual agents to economic events is unlikely to occur simultaneously. In contrast, individual agents are not expected to act uniformly at the same time to a given economic event, rather some agents may be able to response instantaneously, while others will be able to adjust in different time lags. As a result, instead of modeling structural change in trend as an instantaneous trend break, smooth transition models permit the possibility of a smooth transition between different trend paths over time. The different trend paths represent different regimes over time, with the latter attributed to changes or reforms in economic or other general conditions.

In this aspect, Leybourne et al. (1998), hence LNV, consider the following logistic smooth transition models as alternative to difference stationarity,

$$\begin{aligned}
&\text{Model A}: y_t = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + v_t \\
&\text{Model B}: y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + v_t \\
&\text{Model C}: y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 t S_t(\gamma, \tau) + v_t
\end{aligned} \tag{1}$$

where v_t is a zero-mean stationary process and $S_t(\gamma, \tau)$ is the logistic curvilinear function, based on a sample size T,

$$S_t(\gamma, \tau) = \{1 + \exp\{-\gamma(t - \tau T)\}^{-1}, \quad \gamma > 0 \tag{2}$$

that controls the transition between trend paths or regimes.[2] In the logistic function $S_t(\gamma, \tau)$, the parameters of interest are τ and γ, establishing the position and speed of the transition between regimes, respectively. In particular, the location parameter τ determines the timing of the transition midpoint, as for $\gamma > 0$ we have $S_{-\infty}(\gamma, \tau) = 0$, $S_{\infty}(\gamma, \tau) = 1$ and $S_{\tau T}(\gamma, \tau) = 0.5$. The slope parameter γ determines the speed of the transition. For small values of γ, it takes a long period of time for $S_t(\gamma, \tau)$ to traverse the interval (0,1), and in the limiting case with $\gamma = 0$, $S_{\tau T}(\gamma, \tau) = 0.5$ for all t. In contrast, for large values of γ, function $S_t(\gamma, \tau)$ traverses the interval (0, 1) very rapidly, and the case where γ approaches infinity $+\infty$ implies the presence of a threshold model with an abrupt change at time τT.

The parameters α_2 and β_2 determine transition direction in the intercept and trend, respectively. In particular, under the assumption of a stationary v_t process, then in Model A y_t is stationary around a mean which changes from an initial value α_1 to $(\alpha_1 + \alpha_2)$, while Model B is similar and also allows for a fixed trend β_1. In the most general case of Model C, along with the change $(\alpha_1 + \alpha_2)$ in intercept, there is a simultaneous change (of the same speed) in trend, which changes from the initial value β_1 to a final of $(\beta_1 + \beta_2)$. The LNV unit root tests are based on the following set of hypotheses:

$$\begin{aligned}
&\text{Null hypothesis}: y_t = \mu_t, \mu_t = \mu_{t-1} + \epsilon_t,\ \mu_0 \text{ constant} \\
&\text{Alternative hypothesis}: \text{ Models A, B, or C}
\end{aligned}$$

$$\begin{aligned}
&\text{Null hypothesis}: y_t = \mu_t,\ \mu_t = \kappa + \mu_{t-1} + \epsilon_t,\ \mu_0 \text{ constant} \\
&\text{Alternative hypothesis}: \text{Models B or C}
\end{aligned}$$

where ϵ_t is assumed to be a stationary zero mean processes. Then, the test statistics can be calculated via a two-step procedure as follows:

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[2] Upon estimation, time t in the transition function is scaled between 0 and 1 replacing t with t/T and $t - \tau T$ with $t/T - \tau$.

Step 1 : Using a non-linear least squares (NLS) algorithm, estimate only the deterministic component of the corresponding model A, B or C in (1) and compute the NLS residuals $\hat{v}_t$

Step 2 : Compute the ADF statistic, the t-ratio associated with ρ in the ordinary least squares (OLS) regression

$$\Delta \hat{v}_t = \rho \hat{v}_{t-1} + \sum_{i=1}^{k} \delta_i \Delta \hat{v}_{t-i} + \eta_t \tag{3}$$

where k augmentation terms are included to parametrically capture correlation in v_t. This ADF type statistic is denoted as s_α, $s_{\alpha(\beta)}$ or $s_{\alpha\beta}$ when residuals $\hat{v}_t$ are calculated from Model A, B or C in (1), respectively.

Harvey and Mills (2002), hence HM, extend the LNV tests to allow for a second structural change occurring during the observation period of the time series being investigated. Under the alternative, they consider three models which represent stationary processes around two smooth transitions in the linear trend,

$$\begin{aligned}
\text{Model A} : y_t &= \alpha_1 + \alpha_2 S_{1,t}(\gamma_1, \tau_1) + \alpha_3 S_{2,t}(\gamma_2, \tau_2) + v_t \\
\text{Model B} : y_t &= \alpha_1 + \beta_1 t + \alpha_2 S_{1,t}(\gamma_1, \tau_1) + \alpha_3 S_{2,t}(\gamma_2, \tau_2) + v_t \\
\text{Model C} : y_t &= \alpha_1 + \beta_1 t + \alpha_2 S_{1,t}(\gamma_1, \tau_1) + \beta_2 t S_{1,t}(\gamma_1, \tau_1) \\
&\quad + \alpha_3 S_{2,t}(\gamma_2, \tau_2) + \beta_3 t S_{2,t}(\gamma_2, \tau_2) + v_t
\end{aligned} \tag{4}$$

where v_t is again a zero-mean stationary process and

$$S_{it}(\gamma_i, \tau_i) = \{1 + \exp\{-\gamma_i(t - \tau_i T)\}^{-1}, \quad i = 1, 2$$

are logistic smooth transition functions as in (2).

As with the LNV tests, Model A contains no trend and implies a double transition in the mean only, Model B involves a double mean transition, but allows for a fixed slope term β_1, while Model C involves a double transition in both intercept and trend. The speed and midpoints of the two transitions are given by $(\gamma_1, \tau_1 T)$ and $(\gamma_2, \tau_2 T)$, respectively. The tests of the unit root hypothesis against one of the alternatives in (4) are conducted by the same two-step procedure of LNV tests, and the corresponding test statistics are now denoted as $s_{2\alpha}$, $s_{2\alpha(\beta)}$ or $s_{2\alpha\beta}$ for models A, B, C, respectively. Vougas (2006, Table 1, p. 799) and Harvey and Mills (2002, Table 1, p. 677), provide critical values for these tests and for various sample sizes by Monte Carlo simulations. We tabulate the critical values at the 10, 5 and 1 % significance level for a sample size of 250 observations[3] in Table 7.

Applications of the LNV tests include Greenway et al. (1997, 2000) who explore the possibility of smooth transitions in GDP growth for various countries, and Leybourne and Mizen (1999), who examine for smooth transitions in consumer price indices

[3] Our sample consists of 253 quarterly observations on labor force participation rates.

(CPI). Further extensions with respect to non-linear trend stationary alternatives have been considered by Sollis et al. (1999) and Sollis (2005).

3 Empirical application

3.1 US labor supply descriptives

To begin with, we briefly describe the trends in labor supply in the United States since 1948. We employ seasonally adjusted quarterly data from the Bureau of Labor Statistics that cover the period 1948:1 to 2011:1 for both gender- and age-specific labor force participation rates. Data are used in quarterly frequency in several studies that examine the trends in US labor force participation rates (e.g., McEwen et al. 2005), while the use of quarterly data can be further associated to multivariate analysis, for example, VAR models, that employs macroeconomic time series usually found at the same frequency. Foreshadowing our findings, the deterministic transitions in labor force participation rates are gradual and require a great number of quarters to be completed; hence, even if monthly data were employed it would not affect the results with respect to the smooth nature of the transitions. The use of monthly data to jointly inspect the presence of gradual deterministic trends and other non-linearities related to high frequency business cycle factors is left for future research. On the other hand, the use of annual data might raise the question of aggregation issues.

We examine both female and male participation rates by eight age groups (16-over, 16–19, 20–24, 25–34, 35–44, 45–54, 55–64, 65-over). Information regarding summary statistics of the variables employed is given in Table 1. The reported mean levels quantify the movement in labor force participation rates over the sample.

We first focus on female participation rates, which have presented the most outstanding movements in the post-war era. As Fig. 1 illustrates, aggregate female participation rate (16-over) exhibits an almost monotonic and considerable rise after the mid-1960s until the early 1990s. Aggregate participation rate has almost doubled from around 33 % in 1948 to 60 % in 1999. During the 1990s, the growth rate slowed substantially, resulting in a slight decline during the 2000s. The rise in married women's (possibly with small children) proportion in the labor force is considered as an essential factor for the increase in the aggregate participation rate. Female groups aged 20–24, 25–34, and 35–44 display approximately the same trends for the respective time period.

Although young women aged 16–19 have increased their participation rate from the mid-1960s until the mid-1980s, they have reduced it afterwards, resulting in a sharper fall after the early 2000s. For example, participation rate fell from a point of almost 52 % in 2000 at a point of almost 35 % in 2010. The interaction of schooling and work decisions, along with the increase in educational attainment for younger workers during the last decades are considered the key factors for this trend. Another explanation for this downward trend, especially after 2000, is related to the type of jobs that young people undertake when entering the labor market. Generally, younger, less-skilled and less-experienced workers are hired in temporary jobs, characterized by high turnover (i.e., quick hiring and firing or jobs with temporary contracts) that make them extremely sensitive to economic downturns. As a result, the recession of

Table 1 Descriptive statistics

Age group	Mean subsample 1948:1–1952:4	SD	Mean subsample 1977:1–1981:4	SD	Mean subsample 2006:1–2011:1	SD	Min full sample 1948:1–2011:1	Max
Panel A. Females								
16-over	33.81	0.845	50.59	1.348	59.15	0.395	32.20	60.10
16–19	42.04	1.083	52.78	1.344	39.32	3.351	33.90	54.90
20–24	45.55	0.997	68.45	1.195	69.46	0.878	42.70	73.80
25–34	34.27	1.139	63.60	2.611	74.71	0.399	31.60	76.80
35–44	38.86	1.360	63.41	2.724	75.67	0.491	36.60	78.40
45–54	37.72	2.188	58.42	2.044	75.92	0.274	33.70	77.10
55–64	26.57	1.920	41.33	0.574	59.19	0.931	23.50	60.80
65-over	9.28	0.671	8.16	0.254	13.04	0.828	7.10	14.10
Panel B. Males								
16-over	86.40	0.294	77.54	0.381	72.46	0.992	70.40	86.90
16–19	62.84	1.145	60.80	1.256	39.17	3.307	34.30	64.80
20–24	87.12	1.923	85.86	0.509	77.42	2.100	74.00	90.60
25–34	96.45	0.713	95.19	0.222	91.00	1.099	89.10	97.90
35–44	97.76	0.308	95.56	0.230	91.92	0.457	91.10	98.40
45–54	95.86	0.343	91.29	0.264	87.64	0.693	85.90	97.20
55–64	87.72	1.019	72.54	1.208	69.92	0.529	64.70	89.80
65-over	45.40	1.871	19.51	0.828	21.31	0.857	15.20	48.70

In percentage. The reported mean levels (and standard deviations) have been computed for three 5-year subsample periods which correspond to the beginning, the middle, and the end of the available sample and quantify the movement in labor force participation rates over the sample. Minimum and Maximum values are calculated over the entire sample

2000 and the most recent of 2007 may have contributed to the substantial decline of participation rate of youths.

We can further see that growth in participation of women 45–54 and 55–64 began earlier, when comparing to the other age groups. Finally, women aged 55–64 and 65-over tend to have higher participation in recent years. Women aged 55–64 present an ongoing trend since the late 1980s, while women 65-over present a more steep rise from the mid-1990s. The latter present an increase from approximately 9 % in 1999 to approximately 14 % in 2010.

Male participation rates exhibit a completely different trend. As Fig. 2 indicates, participation rate has generally been falling during the last 60 years, both in the aggregate and for most age groups. The rate of the reduction though differs between age groups. For example, the participation rate of young males aged 16–19 fell from around 65 % in 1948 to almost 34 % in 2011, while the decrease for age group 35–44 varies from almost 98 % in 1948 to almost 91% in 2011. The decline in participation rate of young men aged 16–19 also portrays the rise in educational attainment of this group and is more steep during the last 11 years, in step with young women's participation addressed above.

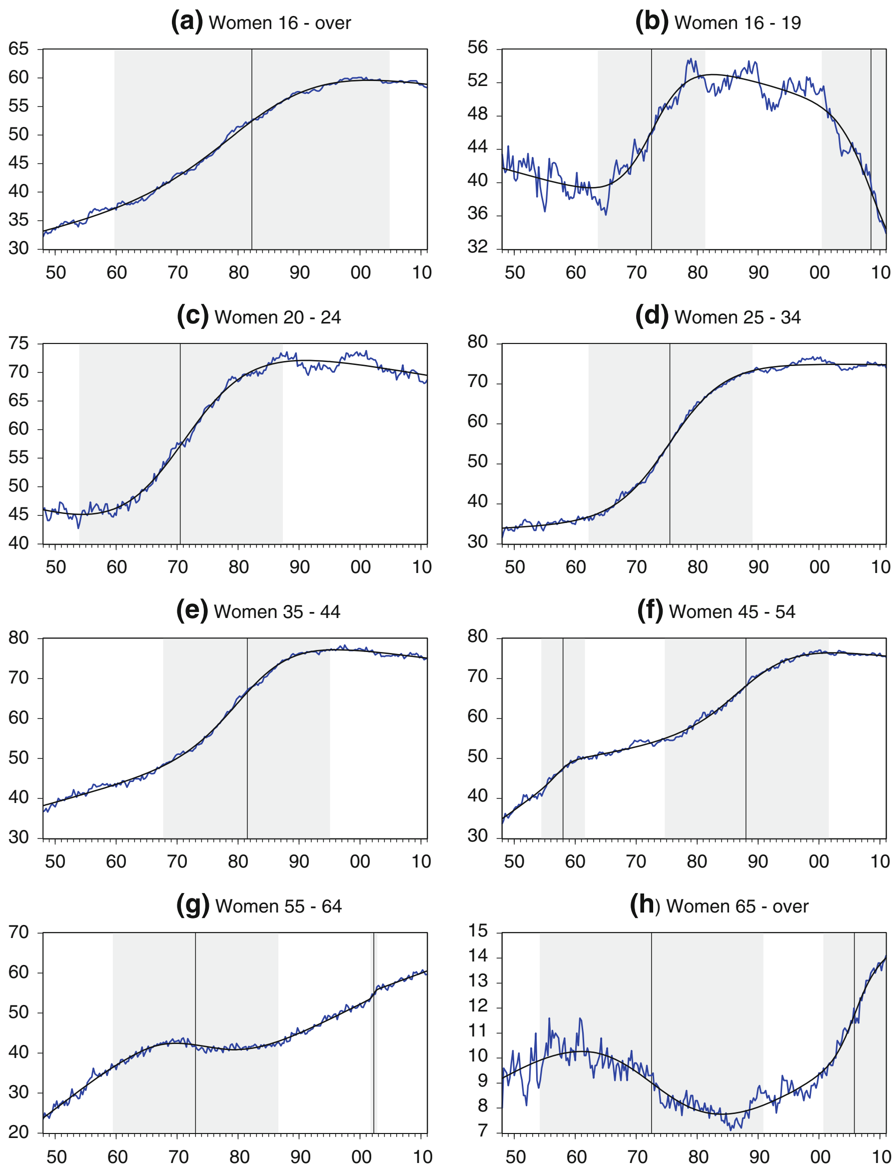

Fig. 1 Female (age-specific) labor force participation rates and fitted smooth transition trends. *Vertical lines* located at estimated midpoints. *Shaded areas* denote 90 % estimated transition range

The reduction in the participation rate of prime-age males, i.e., those aged 25–54, has been attributed to the decline in demand of less-skilled workers that further turned to a steep fall in less-skilled men's employment. A remarkable exception comprises participation rates of age groups 55–64 and 65-over. These groups present a continuous decline, which was stabilized by the late 1980s. This trend is attributed to the

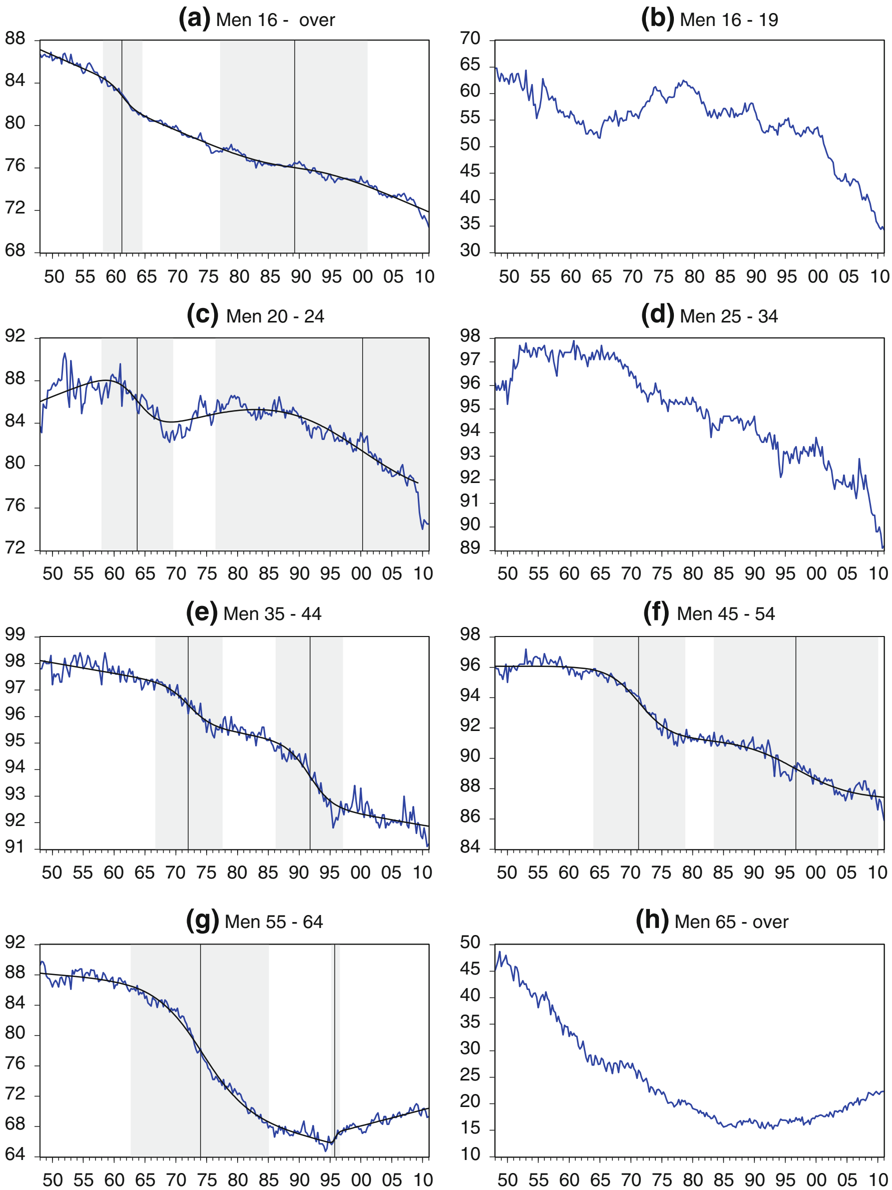

Fig. 2 Male (age-specific) labor force participation rates and fitted smooth transition trends. *Vertical lines* located at estimated midpoints. *Shaded areas* denote 90 % estimated transition range

introduction of Social Security and the retirement of men at younger ages. However, labor force participation rates have tend to rise in recent years. Participation rate of men aged 55–64 has fallen from around 89 % in 1948 to almost 66 % in 1995, and it has been almost 70 % by 2010. Meanwhile, participation rate of men 65-over fell

Table 2 Unit root tests of labor force participation rates (levels)

Panel A. Females			Panel B. Males		
Age group	ADF	DF-GLS	Age group	ADF	DF-GLS
16-over	1.6498	−0.5334	16-over	−1.5776	−1.3585
16–19	0.5426	−0.2765	16–19	−0.6346	−1.0581
20–24	−0.0498	−0.7094	20–24	−0.9518	−0.7443
25–34	−1.1462	−1.5535	25–34	−2.8280	−0.7961
35–44	−1.1438	−1.6973	35–44	−3.5306[b]	−1.2734
45–54	−0.9433	−0.8519	45–54	−3.0733	−1.7451
55–64	−2.0971	−1.4721	55–64	−1.1776	−1.4106
65-over	0.7923	−0.8526	65-over	−0.1112	−0.3584

The unit root tests are applied with a constant plus a time trend (allowing for a deterministic linear trend does not weaken the evidence for non-stationarity). Lag length for the ADF and the DF-GLS tests is selected via the Akaike information criterion

[b] Indicates significance at the 5 % level

from around 45 % in 1948 to almost 15.5 % in 1991, and reached a point of 22.5 % in 2010.

3.2 Unit root and smooth transition unit root tests

We initially apply unit root tests that are widely used in the literature to test whether male and female participation rates are better characterized as unit root non-stationary or as stationary processes around linear deterministic components. In particular, we employ the Augmented Dickey-Fuller (1979, ADF) test and the GLS transformed Dickey-Fuller test (Elliott et al. 1996).[4] Lag length in the ADF and the DF-GLS tests is selected via the Akaike information criterion. Note, however, that the use of the Schwarz and Hannan-Quinn information criteria or using the modified information criteria suggested by Ng and Perron (2001) with the modification proposed by Perron and Qu (2007) point to the same decision of non-stationary participation rates. Based on the results tabulated in Table 2, which reveal a similar behavior of the ADF and the DF-GLS tests, we conclude that all participation rates are non-stationary.

In our case, applying the unit root tests for the first differences of labor participation rates should produce unequivocal evidence of stationarity in order to conclude in favor of I(1) non-stationarity. However, the unit root results on first differences (and then second differences) suggest I(2) behavior for the majority of participation rates.[5] Time series that display non-linear features may distort the behavior of standard unit root/stationarity tests. The non-inclusion of structural breaks or other non-linearities

[4] We have also computed the Phillips and Perron (1988, PP) and the Ng and Perron (2001) MZ_a^{GLS} unit root tests. The results from these tests are not reported (they are available upon request), yet they are in full agreement with the ADF and the DF-GLS test results reported in the paper (with respect to non-stationarity decision). All unit root tests have been carried out in EViews 7.2

[5] For brevity, these results are not reported. They are available upon request.

may bias the results of the unit root and stationarity tests towards non-stationarity, hence manifesting I(2) behavior that is difficult to justify in economic terms. Although more smooth and more slowly changing levels can imply a double unit root, the economic justification of such behavior is not straightforward. For example, the influence of shocks is amplified in time (see Haldrup (1998), p. 604).

As a consequence, we employ the LNV and HM unit root tests to examine the alternative hypothesis of stationary labor force participation rates around smooth transition(s) in linear trend. We estimate the smooth transition ADF test statistics s_{α}, $s_{\alpha(\beta)}$, $s_{\alpha\beta}$ and $s_{2\alpha}$, $s_{2\alpha(\beta)}$, $s_{2\alpha\beta}$ for all age-specific groups. Non-linear model selection was based on a general to specific procedure where the most complex model (HM model C) is estimated and, upon strong convergence, we inspect statistical significance of parameters α_i, β_i, $i = 1, 2, 3$ given that $s_{2\alpha\beta}$ rejects the null of non-stationarity. Heteroskedasticity and autocorrelation consistent standard errors were computed in all cases.

The non-linear least squares (NLS) estimation of (1) and (4) is conducted using the Broyden, Fletcher, Goldfarb, and Shanno (BFGS) optimization algorithm in Ox 5.10. A fine grid search of starting values for γ, τ and γ_1, γ_2, τ_1, τ_2 was considered. In total, 3, 582 regressions were estimated for each series and each test based on (1) while a total of 52, 500 regressions were estimated for each series and each test based on (4). The number of lags included in the unit root tests is calculated by a general to specific procedure where an AR(20) regression model for $\Delta\hat{v}_t$ was estimated for each series and insignificant augmentation terms were excluded. Then, regression (3) with selected augmentations was estimated to compute the test statistics. In order to address heteroskedasticity during the lag augmentation procedure, we employed White (1980) heteroskedastic consistent standard errors. The smooth transition unit root results are presented and discussed in the following section. [6]

3.3 Discussion of the smooth transition unit root test results

3.3.1 Female labor force participation rates

We first focus on female labor force participation rates. For all age-specific groups, we reject the null hypothesis of non-stationary unit root participation rates in favor of the alternative of stationarity in the presence of smooth transition(s) or structural changes in the deterministic component. The rejections occur at the 1 %-level for most participation rates. The estimated midpoints of each transition correspond with the movements of the female participation rates described earlier in the paper.

Table 3 presents the results for female groups with respect to the selected model, the corresponding smooth ADF test statistic and the estimated midpoints and speeds of transition. In addition, Table 4 presents the range of the transitions, that is the 50, 80, and 90 % of the completed transitions centered at the midpoints, and the relative speed or duration of the transitions, measured by the number of quarters required to

[6] All results reported from the smooth transition unit root tests were obtained using programs written by the authors in Ox version 5.10, see Doornik (2007).

Table 3 Smooth transition ADF test results for labor force participation rates

Age group	Test statistic		Lags (k)	Midpoint (τ_1)	Speed (γ_1)	Midpoint (τ_2)	Speed (γ_2)
Panel A. Females							
16-over	$s_{\alpha\beta}$	-5.0548^{b}	6	1982:2	2.395		
16–19	$s_{2\alpha(\beta)}$	-6.7183^{a}	3	1972: 3	6.137	2008:3	6.647
20–24	$s_{\alpha\beta}$	-6.6851^{a}	11	1970:3	3.251		
25–34	$s_{\alpha\beta}$	-6.1930^{a}	7	1975:3	4.001		
35–44	$s_{\alpha\beta}$	-6.4047^{a}	9	1981:3	3.960		
45–54	$s_{2\alpha\beta}$	-7.5329^{a}	5	1958 :1	15.484	1988:1	4.021
55–64	$s_{2\alpha\beta}$	-8.3611^{a}	5	1973:1	4.010	2002:2	89.425
65-over	$s_{2\alpha(\beta)}$	-6.6441^{b}	4	1972:3	2.963	2005:4	10.644
Panel B. Males							
16-over	$s_{2\alpha(\beta)}$	-5.8273^{b}	3	1961:2	16.995	1989:2	4.521
16–19	$s_{2\alpha\beta}$	-5.7851	0				
20–24	$s_{2\alpha(\beta)}$	-6.9885^{a}	9	1963:4	9.253	2000:2	2.192
25–34	–	–					
35–44	$s_{2\alpha(\beta)}$	-11.288^{a}	11	1972:1	10.011	1991:4	9.902
45–54	$s_{2\alpha}$	-6.2707^{a}	10	1971:2	7.260	1996:4	4.051
55–64	$s_{2\alpha\beta}$	-7.6574^{a}	8	1974:1	4.847	1995:4	77.830
65-over	–	–					

Notes The number of lags (k) refers to the total number of lags included in the regression
[a] and [b] denote significance at the 1 and 5 % level respectively. Critical values for LNV tests are reported in Vougas (2006, p. 799) and for HM tests in Harvey and Mills (2002, p.677). We display the relevant critical values in Table 7

traverse 50, 80, and 90 % of the transition path. Figure 1 shows the fitted smooth transition deterministic components.

We find evidence of stationarity around a single smooth transition in intercept and trend for participation rates of women aged 16-over, 20–24, 25–34, and 35–44 (Fig. 1a, c–e). The estimated midpoint of each transition occurs within the time period that corresponds to the substantial rise in female participation rates, and is observed at 1982:2, 1970:3, 1975:3, and 1981:3, respectively. The range and speed of each transition strengthen the notion of gradual or smooth changes in participation rates, since all speeds are relatively small. For example, Table 4 indicates that the 50 % of the transition in group 20–24 takes place within 49 quarters (approximately 12 years), while the 90 % of the transition is completed within 133 quarters (approximately 33 years).

Similarly, the 90 % of the transition of women 25–34 and 35–44 is completed within 107 and 109 quarters (almost 27 years), respectively. Notice that the midpoint of the transition in the aggregate participation rate of women (16-over) occurs later when compared to the other age groups, while ranking the speeds (among all female groups) shows that this group exhibits the slowest transition with the 50 and the 90 % of the transition being completed within 17 and 45 years, respectively.

Table 4 Speed and range of transition, female labor force participation rates

Age group	Range of transition				Speed of transition (number of quarters)		
	90 % transition	80 % transition	50 % transition	Midpoint	90 % transition	80 % transition	50 % transition
16-over	1959:4–2004:4	1965:3–1999:1	1974:1–1990:4	1982:2	180	134	67
16–19	1963:4–1981:2	1966:1–1979:1	1969:2–1975:4	1972:3	70	52	26
	2000:3–2016:4[a]	2002:3–2014:4[a]	2005:3–2011:4	2008:3	65[a]	49*	25
20–24	1954:1–1987:2	1958:2–1983:1	1964:3–1976:4	1970:3	133	99	49
25–34	1962:2–1989:1	1965:3–1985:4	1970:3–1980:4	1975:3	107	81	41
35–44	1967:4–1995:1	1971:2–1991:3	1976:2–1986:3	1981:3	109	81	41
45–54	1954:3–1961:3	1955:3–1960:3	1956:4–1959:2	1958:1	28	20	10
	1974:4–2001:3	1978:2–1998:1	1983:2–1993:1	1988:1	107	79	39
55–64	1959:3–1986:3	1963:1–1983:1	1968:1–1978:1	1973:1	108	80	40
	2001:4–2002:4	2001:4–2002:4	2002:1–2002:3	2002:2	4	4	2
65-over	1954:2–1990:4	1959:1–1986:1	1965:4–1979:2	1972:3	146	108	54
	2000:4–2011:1	2002:1–2009:3	2004:1–2007:4	2005:4	41	30	15

*Denotes estimated transition range that ends after the latest point in the data

On the other hand, participation rates of women aged 16–19, 45–54, 55–64, and 65-over exhibits two structural changes (Fig. 1b, f–h). Among these series, women 16–19 and 65-over involve transitions in the intercept only, while presenting a fixed trend component, whereas women 45–54 and 55–64 present transitions in both intercept and trend. The midpoints of the first transition occur during the 1970s for all groups except for women aged 45–54. This group exhibits an earlier transition, owing to the earlier increase in the participation rate. (Fig. 1f). The speed of this transition is relatively high. The 50 % of the transition is completed within 10 quarters (almost 2.5 years) and the 90 % within 28 quarters. All remaining first transitions are rather smooth. For example, 90 % of the transition of women 65-over requires 146 quarters (almost 36 years) to be completed.

The second structural changes occur after the 2000s. Older women 55–64 and 65-over now adjust more quickly, when comparing to the first transitions. Ranking the speeds reveals that the higher transition occurs for women 55–64, for which the 50 % of the transition is completed within 2 quarters (less than a year) and the 90 % in 4 quarters. In addition, for women aged 65-over the 50 % of the transition occurs within 15 quarters (almost 4 years) and the 90 % is completed within 41 quarters (10 years). On the other hand, women 16–19 and 45–54 exhibit gradual second transitions. Indeed, the second transition of women 16–19 is so gradual, so that the estimated 80 % and the 90 % transition ranges end after the latest point in our data. Notice also that women 45–54 show the earlier second midpoint, in 1988:1.

We now proceed with a closer examination of the estimates of the smooth transition models, with emphasis placed on the changes in the slope of participation rates. In particular, the trend growth rates of the participation series under the former regime (before the first transition) can be calculated from the estimates of β_1, and the augmentation of β_2 (and β_3 in the cases of double transitions) estimates to β_1 gives the trend growth rates in the second and third regime, respectively. Panel A of Table 6 presents smooth transition estimates for female participation rates.

With respect to the initial trend growth rates, all β_1 estimates are positive except for young women 16–19 and 20–24, which exhibit a negative trend growth rate in step with the reduction of participation rates of these age groups related to the increase in school enrollment and education. These groups present a -0.05 and -0.08% reduction per quarter, respectively.

For the remaining groups, we further observe a greater rise in the trend growth rate of participation rates as we move to groups with higher aged components. The β_1 estimate for women aged 25–34 implies a rise in the participation rate on the order of 0.014 % per quarter under the former regime. The corresponding rise for groups 35–44, 45–54, and 55–64 is about 0.103, 0.271, and 0.294 % per quarter, respectively. Thus, women 45–54 and 55–64 exhibit an almost twenty times greater increase in their participation rate with respect to women 25–34, and an almost three times greater increase with respect to women 35–44.

As a result, we conclude that higher-aged female groups present a more decisive (re)-entrance in the labor market, which can be associated with diminished fertility of older-aged women. In other words, women 45–54 and 55–64 which have past the childbearing years appear to strongly participate in the labor market, whereas younger women 25–34 and 35–44 that represent ages of high childbearing show a lower increase in their participation in the labor market. Finally, the older group of women aged 65-over displays a fixed positive trend growth rate of almost 0.04 % per quarter, while the aggregate group presents an increase in the trend growth rate of almost 0.06 % per quarter.

We next examine the results under the second (or third) regimes, after the realization of the transitions. All female groups, except for older women whose participation tends to rise in recent years, exhibit similar reductions in the trend growth rate of participation rate. Women aged 16-over, 25–34 and 35–44 (Fig. 1a, d, e) exhibit reductions in the latter regime, since all β_2 parameters are negative. Adding β_2 to the (positive) estimates β_1 of the first regime, results to a negative slope in the second regime that reflects the slight decline in the participation rates observed after the 2000s.

In particular, the aggregate group exhibits a negative trend growth on the order of -0.04 % per quarter. Women 20–24 (Fig. 1c) also display a negative slope in the latter regime, yet the trend growth rate is now reduced by almost 50 %, implying a slower reduction in the participation rate in recent years. Women 25–34 present a much slower decline, near -0.005 % per quarter, suggesting a constant long-run participation rate for this group. Women 35–44 exhibit a negative trend growth rate on the order of -0.04 % per quarter.

Women aged 45–54 and 55–64 were found to exhibit double smooth transitions. Women 45–54 (Fig. 1f) show a reduction in the trend growth rate in the second regime; however, the augmentation of β_2 to β_1 continues to indicate a positive trend growth rate in the participation rate. However, the second transition leads to a reduction in the change of the participation rate eventually of almost -0.04 % per quarter. On the other hand, as Fig. 1g indicates, though women 55–64 display two sequential reductions in their trend growth rate during the two transitions, their final growth rate remains positive and almost 0.15 % per quarter. All estimates are plausible and consistent with the movements and trends of age-specific female labor force participation rates.

3.3.2 Male labor force participation rates

We now turn to male participation rates. The results now differ on the number of the series that reject the unit root null hypothesis and the number of transitions, which are more than one. We reject the null in favor of the alternative of stationarity around two smooth transitions in linear trend for five age-specific groups, namely 16-over, 20–24, 35–44, 45–54, and 55–64. Most rejections occur at the 1 % level. Men 16–19 constitutes the only group from all participation rates that does not reject the null hypothesis of a unit root, while men 25–34 and 65-over display a different behavior to be discussed later in the section. Panel B of Table 3 and Table 5 present the corresponding results for male participation rates, while Fig. 2 shows the respective fitted smooth transition components.

Among the series that reject the null hypothesis, only men 55–64 (Fig. 2g) involve changes in trend. Men 16-over, 20–24[7] and 35–44 (Fig. 2a, c, e) involve two transitions

[7] Participation rate of men 20–24 dropped rapidly after the 2007 recession (Fig. 2c), and this decrease affected the NLS estimation and the calculation of the non-linear ADF statistic. In order to continue testing for structural changes in this participation rate, we decide to cut the last seven observations.

Table 5 Speed and range of transition, male labor force participation rates

Age group	Range of transition				Speed of transition (number of quarters)		
	90 % Transition	80 % Transition	50 % Transition	Midpoint	90 % Transition	80 % Transition	50 % Transition
16-over	1958 : 2 – 1964:3	1959 : 1 – 1963:4	1960 : 2 – 1962:3	1961:2	25	19	9
	1977 : 2 – 2001 : 1	1980 : 2 – 1998 : 1	1984 : 4 – 1993 : 4	1989:2	95	71	36
16–19	–	–	–	–	–	–	–
20–24	1958 : 1 – 1969 : 3	1959 : 3 – 1968 : 1	1961 : 3 – 1965 : 4	1963:4	46	34	17
	1976 : 3 – 2024 : 1[a]	1982 : 3 – 2018 : 1*	1991 : 2 – 2009 : 1	2000:2			71
25–34	–	–	–	–	–	–	–
35–44	1966 : 4 – 1977 : 3	1968 : 1 – 1976 : 1	1970 : 1 – 1974 : 1	1972:1	43	32	16
	1986 : 2 – 1997 : 1	1987 : 3 – 1995 : 4	1989 : 3 – 1993 : 4	1991:4	43	33	17
45–54	1964 : 1 – 1978 : 4	1965 : 4 – 1977 : 1	1968 : 3 – 1974 : 1	1971:2	59	45	22
	1983 : 3 – 2010 : 1	1987 : 1 – 2006 : 4	1992 : 1 – 2001 : 4	1996:4	106	79	39
55–64	1962 : 4 – 1985 : 1	1965 : 3 – 1982 : 2	1969 : 4 – 1978 : 1	1974:1	89	67	33
	1995 : 2 – 1996 : 3	1995 : 2 – 1996 : 2	1995 : 3 – 1996 : 1	1995:4	6	5	2
65-over	–	–	–	–	–	–	–

*Denotes estimated transition range that ends after the latest point in the data

Table 6 Smooth transition estimates for labor force participation rates

Age group	α_1	α_2	β_1	β_2	α_3	β_3
Panel A. Females						
16-over	32.692	37.046	0.059	−0.101		
16–19	41.788	18.985	−0.051		−18.879	
20–24	45.447	35.653	−0.083	0.037		
25–34	33.830	42.212	0.014	−0.019		
35–44	38.062	49.237	0.103	−0.151		
45–54	34.640	11.873	0.271	−0.205	39.469	−0.106
55–64	23.508	−12.402	0.294	−0.097	13.695	−0.055
65-over	9.3053	−8.3556	0.039		3.2793	
Panel B. Males						
16-over	87.203	−2.0401	−0.063		2.7099	
16–19	–	–	–	–	–	–
20–24	86.019	− 6.9757	0.059		−20.383	
25–34	–	–	–	–	–	–
35–44	98.120	−1.4327	−0.010		−2.2826	
45–54	96.068	−4.7902			−4.0132	
55–64	88.261	−13.436	−0.016	−0.031	−17.939	0.101
65-over	–	–	–	–	–	–

in the intercept around a fixed slope trend and men 45–54 (Fig. 2f) contain no trend and involve transitions in the mean only. Recall that male participation rate has generally been falling for all age groups during the last 60 years, strongly related to the decline in the demand for less-skilled labor since about 1970 and the decline in real wages (of less-skilled workers) occurred at the early 1970s.

The estimated midpoint of the first transitions are in accordance with this fact, since they occur during the 1960s or 1970s. For example, the midpoint is observed at 1972:1, 1971:2, and 1974:1 for men 35–44, 45–54 and 55–64 respectively. Ranking the speeds reveals smooth changes in these transitions. For example, for men 35–44 the 90 % of the transition requires 43 quarters (almost 11 years) to be completed, while the corresponding duration for men 55–64 is even greater, almost 23 years. In terms of the aggregate participation rate (16-over), the transition is more rapid in the sense that the 90 % of the transition requires 25 quarters (almost 6 years) to be completed.

The second structural changes occur after the 1990–91 recession, with the exception of the aggregate group with estimated midpoint a little earlier, in 1989:2. The estimated speeds imply gradual transitions. Only men 55–64 face a rapid second transition in 1995:4, with the 50 % of this transition being completed in 2 quarters and the 90 % in 6 quarters.

The estimates of the smooth transition models in panel 2 of Table 6 confirm the negative growth rate presented in male participation rates. All β_1 parameters are negative, except for young men 20–24 for which it is positive. This group exhibits two reductions in its mean level and constitutes another case for which the second

Table 7 Null critical values for non-linear unit root tests

LNV	0.100	0.05	0.010
Model A	−3.81	−4.12	−4.70
Model B	−4.38	−4.68	−5.25
Model C	−4.68	−4.97	−5.53
HM			
Model A	−4.90	−5.20	−5.80
Model B	−5.44	−5.74	−6.36
Model C	−5.93	−6.21	−6.79

LNV asymptotic critical values are based on a finite sample of 250 observations and the simulations in Vougas (2006) and HM on a finite sample of 200 observations and the simulations in Harvey and Mills (2002)

transition is so gradual that the estimated 80 and 90 % transition range ends after the latest point in the data.

An interesting result pertains with groups 35–44 and 45–54, which constitute a large proportion of prime-age males. Both groups exhibit a further reduction in their mean levels after the 1990s. In particular, men 35–44 exhibit the change in their mean level in the fourth quarter of 1991, and present a fixed negative growth rate on the order of −0.01 % per quarter, while men 45–54 do not display any apparent trend, yet present a further reduction in their mean level in the fourth quarter of 1996. This result can be related to the decline in demand for less-skilled workers discussed earlier. However, as Murphy and Topel (1997) and Juhn and Potter (2006) point out, the unemployment rate for prime-age males did not increase over the 1990s. In contrast, it was declining, implying that prime-age males that were no longer employed did not enter unemployment but rather exited the labor force. This result implies that researchers should examine participation and unemployment behavior together to draw inference on the labor market conditions.

Older workers 55–64 exhibit a further reduction in the (negative) trend growth rate of participation rate after the first quarter of 1974 (β_2 is negative and −0.016 %, and the augmentation of β_2 leads to a reduction of almost −0.05 %), yet the augmentation of β_3 after the second structural change captures the positive trend growth rate that males 55–64 exhibit after 1995 (Fig. 2g). It has since been increasing by almost 0.06 % per quarter.

We finally discuss the results for the three groups which were not found—for different reasons—to exhibit smooth transitions. Young men 16–19 constitute the only group for which the ADF smooth transition test does not reject the unit root hypothesis (the $s_{2\alpha\beta}$ statistic is −5.7851). Data are supportive for the presence of a stochastic trend or an “always changing trend”,[8] where disturbances play a crucial role in the participation behavior. High turnover that characterizes the jobs in which youths are being hired (discussed in section 3.1) justifies this result.

[8] Perron (2006), Sect. 8.5, reviews and discusses unit root versus trend stationarity in the presence of structural change in the trend function.

Men 25–34 and 65-over make up two cases of participation behavior that were not capable of being modeled with smooth transition regressions. Figure 2d indicates that men 25–34, which represent a very active group of the labor market, exhibit a continuous stepwise decline after the late 1960s and appear to be sensitive to business cycle effects, as if they are the more vulnerable group to exit employment if a recession takes place. Limited experience as opposed to older groups could be responsible for this fact. As a result, men 25–34 would be expected to present multiple structural changes rather than at most two, as implied by the smooth transition models we employed. Similarly, participation rate of men 65-over (Fig. 2h) exhibits more than two structural breaks, though not related to business cycle effects. The first two breaks have midpoints in the mid-1960s and mid-1980s and are associated with stabilization of declining participation rates while the third break is identified by the smooth increase after the mid-1990s.

Summing up our results, we find that the movements in most US labor force participation rates are better characterized as stationary fluctuations around a deterministic non-linear component rather than as non-stationary unit root fluctuations that embody the permanence of stochastic shocks. The gradual nature of the deterministic transitions is verified by the endogenously determined transition speeds and the estimated range and duration of transitions that appear to be completed within a large number of quarters (up to four decades for some series). Moreover, evidence of the estimated transition midpoints and directions is in accordance with the trends of the US participation rates, as they have been described earlier in the paper.

In contrast to the results of Gustavsson and Österholm (2006) and Gustavsson and Österholm (2012), our findings suggest that once a more versatile trend function in the unit root procedure is permitted, e.g., deterministic dynamics in the form of smooth transition functions rather than the standard linear trends, then stochastic trends are no longer perceived in labor force participation rates. In addition, while Ozdemir et al. (2012) conclude in favor of stationary labor force participation rates when structural trend breaks are allowed, yet the estimated trend breaks are considered to occur instantaneously. On the contrary, our findings strongly suggest the presence of gradual deterministic structural changes, while the standard case of instantaneous structural change that is being permitted as a limited case in the testing procedure is only observed in two cases (the instantaneous fashion of these transitions is associated with a very large estimated value of γ).

Our approach is more close to those used in Madsen et al. (2008) and Landajo and Presno (2010) in the sense that the possible non-linear nature of labor force participation rates is emphasized in these studies as well. However, the Caner and Hansen (2001) unit root test applied in Madsen et al. (2008) is related to a different form of non-linearity and is based on the use of a threshold autoregressive (TAR) model, in which stochastic trend dynamics depend on labor market conditions. We should note that the trend models employed in our study are deterministic and should be distinguished from threshold autoregressive (TAR) or smooth transition autoregressive (STAR) models, which are non-linear in the autoregressive parameters.

The added advantage of the methodology used in this study is related to its general treatment of deterministic trends which reflect demographic, cultural, and institutional factors that have historically produced gradual, long-term changes in labor force participation rates. Overall, the smooth transition trend specification in the unit root tests of Leybourne et al. (1998) and Harvey and Mills (2002) provides strong evidence (rejections of the null hypothesis occur at the 1 % significance level for most of the series) of stationary labor force participation rates around non-linear gradually evolving trends.

4 Conclusions

In this paper, we have employed smooth transition (STR) trend models to re-investigate the time series behavior of labor force participation rates in the United States. This approach is based on the presumption that long-run movements in participation rates may be more appropriately modeled as smooth transitions between states or regimes rather than as abrupt mean level changes or as a stochastic trend. Based on the aspect of gradual (non-instantaneous) adjustment of labor force participation rates when aggregate behavior of agents (workers) is considered, we have applied unit root tests that allow stationary participation rates around smooth transitions under the alternative. We have examined labor force participation rates by gender- and age-specific groups to take into account workers' heterogeneity in the participation rates.

Based on our findings from the smooth transition unit root tests, this paper suggests that all female and most male participation rates are better characterized as stationary processes that undergo transitional deterministics. The permission of a more flexible trend function when testing for unit roots provides substantial evidence against the null hypothesis of a unit root and the view of an always changing trend in labor force participation rates. Permanent shocks are allowed; however, their frequency is not as high as the I(1) representation suggests, and their effect is gradually embodied in the long-run participation rates.

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References

Bacon DW, Watts DG (1971) Estimating the transition between two intersecting straight lines. Biomertika 58:525–534

Bai J, Perron P (1998) Estimating and testing linear models with multiple structural changes. Econometrica 66:47–78

Benati L (2001) Some empirical evidence on the 'discouraged worker' effect. Econ Lett 70:387–395

Caner M, Hansen BE (2001) Threshold autoregression with a unit root. Econometrica 69(6):1555–1596

Dagsvik JK, Kornstad T, Skjerpen T (2012) Labor force participation and the discouraged worker effect. Empir Econ. doi:10.1007/s00181-012-0598-9

Darby J, Hart RA, Vecchi M (2001) Labour force participation and the business cycle: a comparative analysis of France, Japan, Sweden and the United States. Jpn World Econ 13:113–133

DiCecio R, Engemann MK, Owyang TM, Wheeler HC (2008) Changing trends in the labor force: a survey. Fed Reserve Bank of St. Louis Rev 90(1):47–62

Dickey DA, Fuller W (1979) Distribution of the estimators for autoregressive time series with a unit root. J Am Stat Assoc 74(366):427–431

Doornik JA (2007) Object-oriented matrix programming using ox, 3rd edn. Timberlake Consultants Press and Oxford, London. www.doornik.com.
Elliott G, Rothenberg TJ, Stock JH (1996) Efficient tests for an autoregressive unit root. Econometrica 64:813–836
Gil-Alana LA (2008) Fractional integration and structural breaks at unknown periods of time. J Time Ser Anal 29(1):163–185
Granger CWJ, Teräsvirta T (1993) Modelling nonlinear economic relationships. Oxford University Press, Oxford
Greenway D, Leybourne SJ, Sapsford D (1997) Modelling growth and liberalisation using smooth transition analysis. Econ Inq 35:798–814
Greenway D, Leybourne SJ, Sapsford D (2000) Smooth transitions and GDP growth in the European union. Manch Sch 68(2):145–165
Greenwood J, Seshardi A, Yorukoglou M (2005) Engines of liberation. Rev Econ Stud 72(1):109–133
Gustavsson M, Österholm P (2006) The informational value of unemployment statistics: a note on the time series properties of participation rates. Econ Lett 92(3):428–433
Gustavsson M, Österholm P (2010) The presence of unemployment hysteresis in the OECD: what can we learn from out-of-sample forecasts? Empir Econ 38:779–792
Gustavsson M, Österholm P (2012) Labor-force participation rates and the informational value of unemployment rates: evidence from disaggregated US data. Econ Lett 116:408–410
Haldrup N (1998) An econometric analysis of I(2) variables. J Econ Surv 12(5):595–650
Harvey DI, Mills TC (2002) Unit roots and double smooth transitions. J Appl Stat 29:675–683
Hotchkiss JL, Robertson JC (2012) Asymmetric labour force participation decisions. Appl Econ 44(16):2065–2073
Im KS, Pesaran MH, Shin Y (2003) Testing for unit roots in heterogeneous panels. J Econom 115(1):53–74
Jacobsen JP (1999) Labor force participation. Q Rev Econ Financ 39:597–610
Johansen S (1988) Statistical analysis of cointegration vectors. J Econ Dyn Control 12:231–254
Juhn C (1992) Decline of male labor force participation: the role of declining market opportunities. Q J Econ 107(1):79–121
Juhn C, Potter S (2006) Changes in labor force participation in the United States. J Econ Perspect 20(3): 27–46
Kapetanios G, Shin Y, Snell A (2003) Testing for a unit root in the nonlinear STAR framework. J Econom 112:359–379
Killingsworth RM, Heckman JJ (1986) Female labor supply: a survey. In: Ashenfelter O, Layard R (eds) Handbook of labor economics, vol 1. Harvard University Press, Cambridge, pp 103–204
Kwiatkowski D, Phillips PCB, Schmidt P, Yongcheol S (1992) Testing the null hypothesis of stationarity against the alternative of a unit root. J Econom 54:159–178
Landajo M, Presno MJ (2010) Nonparametric pseudo-Lagrange multiplier stationarity testing. MPRA Paper No. 25659, http://mpra.ub.uni-muenchen.de/25659/
Lee J, Strazicich MC (2003) Minimum lagrange multiplier unit root test with two structural breaks. Rev Econ Stat 85(4):1082–1089
Lee J, Strazicich MC (2004) Minimum LM unit root test with one structural break. Working paper No 04–17 Department of Economics, Appalachian State University.
Leybourne SJ, Mizen P (1999) Understanding the disinflations in Australia, Canada and New Zealand using evidence from smooth transition analysis. J Int Money Finance 18:799–816
Leybourne SJ, Newbold P, Vougas D (1998) Unit roots and smooth transition analysis. J Time Ser Anal 19(1):83–97
Lin CJ, Teräsvirta T (1994) Testing the constancy of regression parameters against continuous structural change. J Econom 62:211–228
Madsen J, Mishra V, Smyth R (2008) Are labour force participation rates non-stationary? Evidence from 130 years for G7 countries. Aust Econ Pap June, 66–189.
McEwen H, Orrenius R, Wynne M (2005) Opting out of work: what's behind the decline in labor force participation?. Fed Reserve Bank of Dallas, Issue 6
Michael RT (1985) Consequences of the rise in female labor force participation rates: questions and robes. Part 2: trends in women's work, education, and family, building. J Labor Econ 3(1):S117–S146
Murphy KM, Topel R (1997) Unemployment and nonemployment. Am Econ Rev 87(2):295–300
Ng S, Perron P (2001) Lag length selection and the construction of unit root tests with good size and power. Econometrica 69:1519–1554

Ozdemir ZA, Balcilar M, Tansel A (2012) International labour force participation rates by gender: unit root or structural breaks? Bull Econ Res. doi:10.1111/j.1467-8586.2011.00419.x
Pencavel H (1986) Labor supply of men: a survey. Chap. 1. In: Ashenfelter O, Layard R (eds) Handbook of labor economics, vol 1. Elsevier, North-Holland, pp 3–102
Peracchi F, Welch F (1994) Trends in labor force transitions of older men and women. J Labor Econ 12(2):210–242
Perron P (2006) Dealing with Structural Breaks. In: Patterson K, Mills TC (eds) Palgrave Handbook of Econometrics, Econometric Theory, vol 1. Palgrave Macmillan, Basingstoke, pp 278–352
Perron P, Qu Z (2007) A simple modification to improve the finite sample properties of Ng and Perron's unit root tests. Econ Lett 94:12–19
Phillips PCB, Perron P (1988) Testing for unit roots in time series regression. Biometrika 75:423–470
Robinson PM (1994) Efficient tests of nonstationary hypotheses. J Am Stat Assoc 89:1420–1437
Said SE, Dickey DA (1984) Testing for unit roots in autoregressive moving average models of unknown order. Biometrika 71:599–607
Smith PJ, Ward M (1985) Time series changes in the female labor force. J Labor Econ 3(2):59–90
Sollis R (2005) Evidence on purchasing power parity from univariate models: the case of smooth transition trend-stationarity. J Appl Econom 20:79–98
Sollis R, Leybourne SJ, Newbold P (1999) Unit roots and asymmetric smooth transitions. J Time Ser Anal 20(6):671–677
Van Zandweghe W (2012) Interpreting the recent decline in labor force participation. Fed Reserve Bank of Kansas City, Econ Rev, First Quarter, 5–34
Vougas DV (2006) On unit root testing with smooth transitions. Comput Stat Data Anal 51: 797–800
White H (1980) A heteroskedasticity-consistent covariance matrix and a direct test for heteroskedasticity. Econometrica 48:817–838