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Links between Labor Supply and Unemployment: Theory and Empirics

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Keywords: Unemployment, Labor market participation, Female participation

JEL Classification: J1, J6, E24

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Abstract

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1 Introduction

The objective of this paper is to investigate the links between aggregate unemployment and labor market participation, and notably causalities and the empirical problems with testing those links. It uses the insights from a macroeconomic model of labor market participation to illustrate the complex interactions between participation and unemployment and propose an empirical strategy, based on structural VARs, to uncover those links.

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For a long time, labor supply shifts have been absent from empirical works aiming at uncovering the determinants of unemployment. Recent work by Nickell et al. (2005) focus instead on labor market institutions. Blanchard and Wolfers (2000) focus on the interaction between institutions and various macroeconomic shocks but not labor supply shocks. However, in the pace of two decades, the share of women in the active population has risen from 35% to 45% or more in most OECD economies ; the age composition has also dramatically changed, with a younger population in the 1970's and 80's and an ageing one in the 1990's.

Recent works have shown that labor supply has a good potential at explaining rising wage inequality in the US (Topel 1994a and b, Topel 1997, Kim and Juhn 2000, Acemoglu 1999). Such issues had been investigated earlier (Hamermesh and Grant 1979, Grant and Hamermesh 1982): estimate of production functions with different demographic groups show that there is no segmentation of the labor force: an increase in the relative size of a group generates a negative effects on the wage of q-substitute groups and a positive on the wage of q-complement groups[1]. Is it possible that labor supply factors, causing more wage inequality, can also cause more unemployment? This is at least plausible if rigidities in the bottom of the wage distribution prevent downwards wage adjustments.

However, the main difficulty in carrying time-series tests of the causal links between labor supply and unemployment is an endogeneity issue, as unemployment might discourage female participation (as fewer jobs are available). In contrast, added workers effects, a mechanism through which a spouse may decide to participate to the labour market if the primary wage earner is facing unemployment risk, may generate a positive link between unemployment and participation. There are broadly speaking two empirical strategies to deal with these endogeneity issues. A first one, in the spirit of cross-country studies of unemployment (Nickell et al. 2005, Nunziata 2003) is to regress unemployment in panel on various institutions and female participation using with a set of relevant instruments, such as divorce rate or the share of women in national parliaments. Recent empirical works have recently estimated the determinants of participation (Pissarides et al. 2005, Genre et al. 2003, Jaumotte 2004) ; using those studies to implement an instrumental variable strategy is a promising research area. An alternative strategy, explored here, is to have recourse to standard time-series techniques to uncover the double causality between participation and unemployment. Notably, structural VAR techniques can illustrate some of the relationships at work.

In **Section 2**, we expose some recent theoretical advances on the determinants of participation and unemployment in a frictional market, based on Garibaldi and Wasmer (2005). We briefly summarized the findings of the model, and then introduce a few useful additions in our context: the participation rate is made explicit, we develop some dynamics aspects and introduce a calibra-

[1] Groups i and j are q-substitute (resp. q-complement) when $\partial \ln w_i / \partial \ln E_j < 0$ (resp. >0) where E_j is the employment of group j and w_i the wage of group i (see e.g. Hamermesh 1986).

tion of Continental European economies. In **Section 3**, we also discuss various causalities between unemployment and participation on a formal basis, which was not in Garibaldi and Wasmer (2005).

This discussion is the basis for the development of an empirical methodology in **Section 4**. It justifies the use of a SVAR approach to deal with bilateral causalities. The model also has a second goal: it is to derive a proper identification of shocks in the empirical part. Generally speaking, our model would potentially offer two possibilities: the first one is to use the long-run relationship between variables to obtain a *long-run* identification of shocks à la Blanchard and Quah (1989). The second one is to calibrate the model for the US and a few Continental European countries to fit with participation rates and unemployment, so as to derive a *short-run* identification of shocks.

We have to chose one against the other, and our choice is exposed in that section too. The long-run identification could indeed be derived from the model, where a permanent shock on the distribution of utility of leisure has a long-run impact on both participation and on the level of unemployment. This arises because leisure has an impact on wages and thus on job creation and quit decisions. This identification would imply that participation shocks have a permanent effect on unemployment. Even though this makes sense, it is as odd with several macroeconomic models where participation shocks are neutral, because they do not imply a compositional change of the labour force. The issue of long-run neutral labor supply shocks in macroeconomic models was discussed at length in Wasmer (1997) and some key aspects of the discussion will be summarized in that section. In contrast, short-run identification is more natural in a flow model: a shock on labor supply has a short-run impact on unemployment, as time to search for a job is strictly positive.

Section 5 displays the results. In all countries, participation shocks have had a moderate short and medium run impact on unemployment rates, while unemployment shocks reduce participation during a few years. A variance decomposition exercise indicates that, in Continental Europe, participation is driven in the short run by unemployment shocks, while in the US, it is driven by participation shocks, interpreted as demography or immigration. Short-run unemployment is primarily driven by unemployment shocks in the US, but in Continental Europe, participation shocks drive the short-run evolution of unemployment instead. This result is obtained in a VAR with a linear trend: it does not imply that participation is the major cause of rising unemployment in Europe, but rather that employment has not positively co-moved with participation in these countries. **Section 6** concludes.

2 A theory of labor supply in imperfect labour markets

It is not easy to build a theoretical model determining both unemployment and endogenous participation to the labor market. In general, models of unemployment have inelastic labor supply, while models of choice of hours have perfect labor markets instead. Implicitly, the latter class of models assume that unemployment and non-participation do not need to be disentangled. Nevertheless, the empirical literature has recognized for long that in fact, non-participation was distinct from unemployment: Flinn and Heckman (1983) have shown that "unemployed" and "out of the labor force" are behaviorally distinct states. Notably, the determinants of the transition rate from unemployment to employment differ largely from those of the transition rate from out of the labor force to employment (table 1, p. 36). Recently, Jones and Ridell (1999) have shown on Canadian data that the frontier between the two concepts was however difficult to draw, as there exists a significant margin of workers willing to have a job but not actively searching. These are discouraged workers, or marginally attached workers.

There are few models where one can straightforward analyze the joint relations between unemployment and participation to the labor market and at the same time account for the imprecision of the border between the different labour market states. We simplify here the exposition of the model by Garibaldi and Wasmer (2005), GW hereafter.[2] As explained above, the benchmark model proposed here serves two goals. First, to discuss various causalities between unemployment and participation on a sound basis, and second, to help deriving a short-run identification of supply shocks. For that, we will calibrate the model to Europe and the US using institutional parameters such as the level of unemployment benefits and taxes on wages.

2.1 Setup

In this model, workers are ex-ante homogenous (the assumption is relaxed in the discussion of next section) but ex-post heterogenous. Notably, they have a time-varying stochastic utility of non-participation, denoted by θ, which evolves as a Poisson process with parameter μ drawing in a distribution with c.d.f. $F(.)$. The market productivity is y and is first assumed to be constant. Let us denote by θ^q the reservation utility of leisure which makes workers indifferent between quitting and staying on the job, and by θ^ν the reservation utility of leisure which makes workers indifferent between entering the labor market or staying out of labor market participation.

Wages are assumed to be bargained between the employer and the employee: workers's share in the negotiation in denoted by β with $0 \leq \beta \leq 1$. There is a tax τ on wages and workers derive utility during unemployment denoted by b. Next, in a world with search frictions, the job seekers are recruited at a Poisson rate α, made endogenous later on. When $\alpha < +\infty$, one can show that θ^q is larger than θ^ν: quitting is more difficult than entering in the labor market,

[2] In that paper, we focussed on macroeconomic and calibration issues on the US economies and its labor flows in a steady-state. Here, we will develop and simplify the calculation of the participation rate, offer a simple comparative quantitative exercise of Europe and the US, and develop a discussion of short-run dynamics of the model in response to a marginal supply shock.

because of a loss of search capital and surplus on the job. The labor supply decisions are thus summarized by these two cut-off points, and accordingly we can define

$$\begin{aligned} q &= \mu[1 - F(\theta^q)], \\ \varepsilon &= \mu F(\theta^\nu), \end{aligned}$$

as respectively the quit rate to inactivity of employed workers and the entry rate into the labor market of non-participants to the labor market. We established in GW (2005) that such reservation rules exist and satisfy two equations,

$$\text{Workers' Entry:} \qquad \frac{\theta^\nu - b}{\alpha} = \beta \frac{\theta^q - \theta^\nu}{r + \mu + \delta}, \tag{1}$$

$$\text{Workers' Quit} \quad : \quad \theta^q = y(1-\tau) + \frac{\mu}{r+\mu+\delta} \int_{\theta^\nu}^{\theta^q} F(\theta) d\theta. \tag{2}$$

A simplified proof is derived in Appendix A. The quantity $\beta \frac{\theta^q - \theta^\nu}{r+\mu+\delta}$ is the surplus of a job for a marginal worker with utility of leisure θ^ν.

These two equations have opposite slopes in the plan (θ^ν, θ^q). The first one is upward sloping and states that the opportunity cost of participating for the marginally indifferent worker is the expected forgone value of leisure as a non-participant θ^ν/α while the gain is $b/\alpha + \beta \frac{\theta^q - \theta^\nu}{r+\mu+\delta}$. This implies that θ^ν is increasing in θ^q as θ^q raises the surplus. It is also increasing in the exit rate of unemployment α: this is a crucial link as it shows that at the entry margin, participation to the labor market is reduced by a lower exit rate of job finding.

The second equation implies that θ^q is decreasing with θ^ν, which can be seen from a straightforward differentiation: as θ^ν is larger, the surplus of staying on-the-job is lower, and the gap between θ^q and y becomes smaller. This gap is a hoarding term: the marginal quitter has home production above its wage, because she looses her search capital, i.e. the value of past search efforts.

2.2 Definition of stocks

Population size is normalized to one, and we have the following partition of the population:

- for $\theta < \theta^\nu$, there are only labor market participants, both employed and actively job seeker, i.e. unemployed according to the ILO.
- for $\theta > \theta^q$, there are only non-participants to the labor market.
- for completeness, between the two margins θ^ν and θ^q, one finds the marginally attached workers. If they hold a job they don't quit. Out of job, they don't want to search for a new job. They account for Jones and Ridell (1999) marginally attached.

Let us denote by u_{ILO} the ILO definition of the unemployment rate (only counting workers below θ^ν), and by p_{ILO} the implied definition of participation. Calculations of the steady-state stocks in GW give the unemployment rate, while the Appendix A gives the participation rate which was not in GW.

$$\begin{aligned} u_{ILO} &= \frac{\delta+q}{\delta+q+\alpha}, & (3) \\ p_{ILO} &= \frac{\varepsilon}{\varepsilon+q+\Delta}. & (4) \end{aligned}$$

with $\Delta = \mu\left[F(\theta^q) - F(\theta^\nu)\right]\left(\frac{\delta}{\delta+\mu} + \frac{\mu}{\delta+\mu}\frac{\delta+q}{\delta+q+\alpha}\right) > 0$ is proportional to the number of marginally attached workers $F(\theta^q) - F(\theta^\nu)$.[3]

We have therefore a first set of results. In the plane (θ^q, θ^ν), an increase in α shifts the entry curve upward and leaves the quit curve unchanged. It results an increase in θ^ν, and a decrease in θ^q, thus an increase in ε but also an increase in q. The result on unemployment is a priori ambiguous but one can show that the net effect is always a reduction in unemployment. The result on participation is ambiguous too, and can be either positive or negative, depending on whether workers quit more than they enter.

2.3 General equilibrium

A third equation determines the job finding rate α. As is standard in the search-matching literature, we assume an underlying matching technology associating vacancies and unemployed workers. Let ϕ be the ratio of vacancies to unemployment. We set $\alpha(\phi) = \phi^{1-\eta}$ where $0 < \eta < 1$, and let $\alpha(\phi)/\phi$ be the Poisson rate at which firms recruit workers. A free-entry condition of firms states that expected search costs must be equal to their share of the surplus. Let γ be the flow vacancy cost of firms, then,

$$\gamma(1-\tau)\frac{\phi}{\alpha(\phi)} = (1-\beta)\frac{\theta^q - \theta^\nu}{r+\mu+\delta} \qquad (5)$$

Vacancy costs γ multiplied by $\phi/\alpha(\phi)$ represent the expected value of recruitment costs.

2.4 Calibration

We now proceed with a calibration of two stylized economies, Europe and the US. The two economies with different payroll taxes and replacement ratio of unemployment benefits, aiming at representing Europe and the US. Table 1 gives the parameter values and the equilibrium.

[3] An issue is whether the latter population should be classified as active or inactive workers. ILO is every clear about it: they should be considered as inactive workers. A less strict definition of participation, that would include the marginally attached workers, actually leads to a simpler expression for the participation rate $\varepsilon/(\varepsilon+q) > p_{ILO}$, as explained in Appendix A.

	Europe	US
r: discount rate	0.005	
μ: arrival rate of shocks to leisure	0.06	
δ: exogenous job destruction rate	0.011	
y: market productivity	2.0	
$Support$: upper limit of leisure shocks	1.9	
β: bargaining power of workers	0.5	
η: elasticity of matching function	0.5	
γ: flow recruitment cost of firms	3	
τ: payroll taxes	0.4	0.3
b/y: unemployment benefits	0.5	0.25
Results		
θ^{ν}: entry cut-off point	1.16	1.31
θ^{q}: quit cut-off point	1.24	1.51
ϕ: labor market tightness	0.087	0.387
u_{ILO}: ILO unemployment rate	0.103	0.040
p_{ILO}: ILO participation rate	0.597	0.714
α: job finding rate	0.296	0.622
$q = \mu[1 - F(\theta^{q})]$: quit rate to inactivity	0.023	0.015
$\varepsilon = \mu F(\theta^{\nu})$: entry rate into labor force	0.035	0.039

Table 1: Calibration statistics

All parameters, and notably productivity per worker y and the frequency μ and distribution of family shocks $f(.)$, are chosen so as to be identical for Europe and the US. Parameter differing across the two economy are the institutional parameters: the ratio of unemployment benefits to productivity reproduces Cole and Rogerson's (1998) and Mortensen and Pissarides (1994) values for the US, while I chose a conventional value for Europe. Payroll taxes are equal to 0.4 and 0.3 respectively. There were a few differences between this calibration and the one in GW: if μ, r, β, η are identical to GW and have conventional values, we chose here a uniform instead of a beta distribution, which implies slight changes in productivity parameters, matching efficiency, recruitment costs γ and exogenous job destruction rate ($\delta = 0.011$ here and 0.01 in GW). We also have added taxes and offer a calibration of Europe where in GW we focussed on the US only.

With these parameters, we obtain the following value of the unemployment rate (4% in the US and 10.3% in Europe), corresponding to the average over the 1990's. One could argue that the latter is a bit high if compared to the average in the whole European Union, but he have in mind the three largest countries in continental Europe (Germany, France, and Italy) for which the average unemployment is around 10. Participation rates (in the ILO sense) correspond also well to average participation rates, 60% in Europe and 70% in the US.

As regards to flows in and out of activity, we only consider outflows from employment (given that inflows into the labor force are then determined by the stocks and the outflows). The comparability of data across country is problematic: in the US, monthly flow are aggregated into quarterly flows, while in Europe, LFS data are based on yearly transitions, with typically a lot of infra-yearly transitions problems. We match flows from employment to unemployment relatively well, and tend to underestimate flows to inactivity in the US with a better fit in Europe. Appendix B develops the discussion further, and notably insist on data comparability issues.

2.5 Short-run dynamics: a special case

A full development of the dynamics in this model is difficult, because we need to keep track the time evolution of both all stocks and more difficult, of the time-varying distribution of θ within each stock, notably to determine the dynamic of labor demand. This would require a development of the model that is beyond the scope of the paper. Nevertheless, the short-run dynamics is useful to implement a short-run identification. We will thus proceed as follows: first, consider a steady-state as defined in the calibration above. To this steady state, assume that an infinitesimal number of workers enters the labor force. Let us denote this number by ξ^p for reasons clarified later on. This is a non-permanent increase in labor supply. What is the dynamic evolution of the stock of unemployed workers over a time horizon T?

In a constant labor supply world, the survival rate after s units of time following the shock of these ξ^p additional participants is $e^{-\alpha s}$ where α is the outflow rate. But this is not the end of the story: here, those additional unemployed workers remain in the labor force only as long as their value of home productivity remains below θ^ν: so, with intensity $\mu(1 - F(\theta^\nu))$, a fraction of the unemployed also leave the unemployment pool.

Ignoring multiple transitions –that would lead to newly employed loosing their job and add up to the unemployed pool and quitters to renter the labor force again–, we thus obtain a formula for the increase in unemployment:

Result 1. The average number of additional unemployed workers over a period of length T due to a one-shot increase in participation of ξ^p is

$$\xi^p \left(\frac{1}{T} \int_0^T e^{-\widetilde{\alpha}} ds \right) = \xi^p \left(\frac{1 - e^{-\widetilde{\alpha} T}}{\widetilde{\alpha} T} \right) \tag{6}$$

where $\widetilde{\alpha} = \alpha + \mu(1 - F(\theta^\nu))$. In the US, $\mu(1 - F(\theta^\nu)) = 0.0136$ and in Europe, $\mu(1 - F(\theta^\nu)) = 0.0146$ so that the hazard rate is about 5% larger than α in Europe and 2% larger than α in the US.

3 The causal links from participation to unemployment

3.1 Partial equilibrium: exogenous variation in α

Beyond the short-run, we can now use these two benchmark calibrations to understand the long-run co-variations between unemployment and participation. One major parameter affecting participation is the job finding rate of workers α. Variations in α would also affect participation decisions, θ^{ν} and θ^{q}, and thus participation. The first exercise we carry out here is to impose exogenous variations in α around its general equilibrium value denoted by α^{*}. We indeed solve for the partial equilibrium values of θ^{ν} and θ^{q} thanks to the quit and entry margins, in a range $[\alpha^{*}/2\ ;\ 3/2\alpha^{*}]$. One will thus see how much the degree of frictions affect the entry and the exit decisions in the two economies calibrated above. Figure 1 represents the variation of α around $\alpha^{*} = 0.622$ for the US. It shows that, the higher α, i.e. the more efficient labor market from the perspective of workers, the higher the entry rate and also the quit rate. The intuition for a higher quit rate is easy: with higher α, the hoarding behavior of workers is reduced as labor market is fluid. Overall, the quit effect dominates over the entry effect and both unemployment and participation are reduced: here, $cov(u_{ILO}, p_{ILO}) > 0$. In contrast, in a EU-type economy, represented in Figure 2, where taxes and replacement ratio are higher, one observes that the entry and the quit rates still rise with α, but that the entry effect now dominates: here $cov(u_{ILO}, p_{ILO}) < 0$. In other words, in a low participation country, participation is raised by a more efficient labor market, while in a high participation country, participation can be reduced by a more efficient labor market. This suggests that, if the effect of labor market conditions on labor supply decisions of entry and exit are never ambiguous, they also have the opposite consequences on stocks of participation and thus ambiguous effects.

3.2 General equilibrium variations in α

We can now investigate whether there are positive or negative covariations between unemployment and participation when changes in α are driven by changes in productivity or benefits. We do not report simulations of the general equilibrium impact of a change in productivity y but briefly described them: in both a European-type and US-type economy, an increase in y always reduces quits and raise entries. This simultaneously increase participation and reduces unemployment.

The general equilibrium impact of unemployment benefits is as follows. When the replacement ratio goes from 0.5 to 0.25, the job finding rate α increases from 0.4 to 0.63. The intuition is simple. As b declines, wages pressure is lower, so are wages, and jobs become generate higher profits for firms. This shifts the entry of firms up, and raise vacancy openings and thus the job finding rate α. Unemployment thus declines. On participation margins (quits and entries), we have the following results: the decline in b and the resulting higher

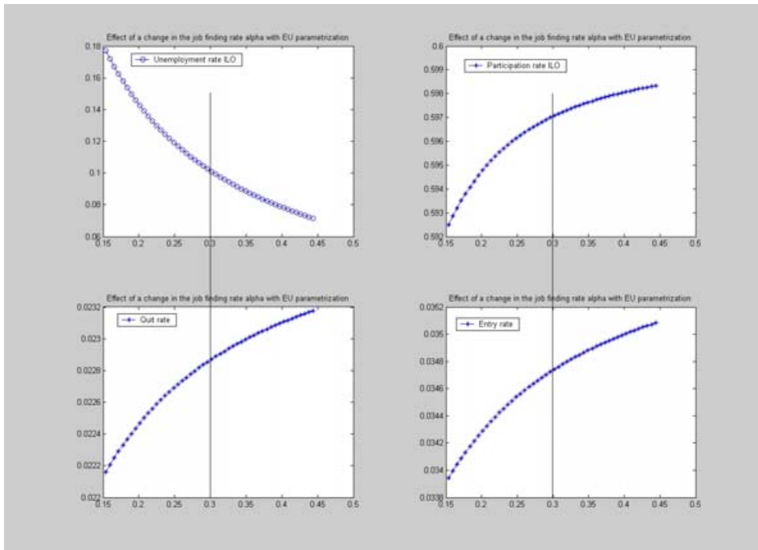


Figure 1: **Partial equilibrium**. United States. X-axis is α. Impact of an exogenous change in α on unemployment, participation, quit and entry. Vertical line is the general equilibrium value of the job finding rate. $\tau = 0.3$, $b/y = 0.25$.

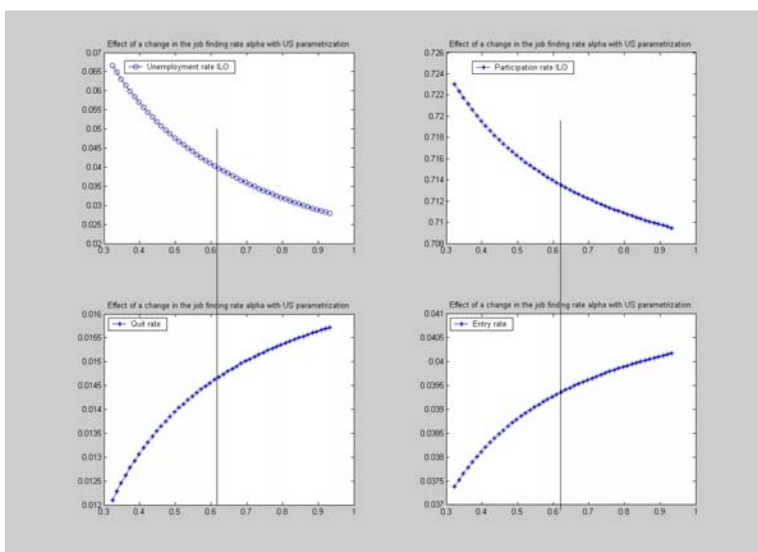


Figure 2: **Partial equilibrium**: Europe. X-axis is α. Impact of an exogenous change in α on unemployment, participation, quit and entry. Vertical line is the general equilibrium value of the job finding rate. $\tau = 0.4$, $b/y = 0.5$.

α leads to a decline of quits because of the reduction of the hoarding effect: in a more efficient labour market, entry and exit decisions (θ^q and θ^ν are closer to each other) and so the gap between wage and reservation home production. On the other hand, lower b leads to a negative effect on entries into the (as participation yields higher income from benefits) and quit since the value of a job relative to non-employment decreases. Overall, participation declines because the disincentive effect on quits dominates.

3.3 Additional effects

3.3.1 Aggregate effects

We first keep the assumption of homogenous labor but relax the assumption that y is constant and exogenous. It could be that the level of aggregate employment has a negative impact on the marginal revenue of the firm. Such would be the case if, along a demand curve, a rise in quantity produced lead to a reduction in prices of goods. In a reduced form, one can assume that $y = F(E)$ where E is aggregate employment with $F' > 0$, and $F'' < 0$.[4] Thus, given that $E = p(1 - u_{ILO})$, participation has a negative impact on y ceteris paribus.[5]

Such a phenomenon is plausible in the short-run when the stock of capital is fixed. In the long-run, the assumption of constant returns to scale is more appealing. This reasoning suggest the existence of a link between the first difference of participation and the first difference of unemployment, which is negative along this productivity (labor demand) curve.

Result 2. In the long-run, one should expect no link between participation and unemployment, because of the homogeneity of labor supply.

3.3.2 Distributional effects

However, changes in labor supply are not homogenous. Assume that there are two segments of the labor force, say female workers and workers substitute to women (young workers, unskilled prime age men[6]). An increase in the level of

[4] This effect can easily be generated in assuming that all goods produced by workers are intermediate goods sold in a competitive market to a final good sector with decreasing returns to scale. The price of each intermediate good thus declines with employment. As a consequence, the marginal revenue of labor declines with aggregate employment. Formally, this is equivalent to having y declining with E.

[5] As pointed out by a referee, the conventional bargaining solution no longer applies when there are decreasing returns to scale, as first developped by Stole and Zwiebel (1996). They call the new solution for wages intrafirm bargaining. Cahuc, Marque and Wasmer (2004) showed that in search context, the marginal product is weighted by a term depending on bargaining power and technological parameters such as the gap to constant returns to scale (one labour type) or substituability between groups of labour (several labour types), resulting into over or under employment of some groups compared to the conventional wage solution. Unreported calculations however indicate that qualitative results are the same as in the benchmark search models, the difference beeing only quantitative.

[6] See the early empirical estimates of production function (Grant and Hamermesh 1982, Hamermesh and Grant 1979, Becker 1983) who find this pattern of substitutability.

employment of female workers (denote them with superscript f) may be detrimental to labor market prospects of those substitute workers (denote them by superscript σ). Since by definition, $\partial^2 F/\partial e^\sigma \partial e^f$ is negative, $y^\sigma(e^f)$ is negatively sloped, one may expect a positive relation between unemployment of group σ and female participation. Works by Topel (1994a and b, 1997) indicate that q-substituabilty exists between unskilled men (below 33th centile) and unskilled women (below median), but also between high skill women (above median) and low skill men.

Along similar lines but in imperfect labor markets, one can imagine in a reduced form the existence of crowding-out in the labor market for substitute workers. Then, $\alpha^\sigma(p^f)$ could be negatively sloped if employers prefer women to young workers at the hiring level, independently of (or more precisely, in addition to) the productivity effects described above. Similarly, rising female participation may lead to displacement of σ-workers.

Result 3. In the short-run, a change in the composition of the labor force due higher participation of a group may raise unemployment above the level implied by identity 6.

Denote by Λ this additional fraction, which depends on the second cross-derivative of production with respect to labor groups and is thus unobserved. The knowledge of Λ or at least a prior on its value is required in order to estimate the impact of rising female participation. In the text below, we will describe a procedure to infer about its value.

Whether Result 2 or Result 3 holds depends on the type of the shock and can be tested.

4 Empirical strategy

4.1 Separating u(p) and p(u)

The toolbox of econometric analysis allows for several strategies, each of which being specific to the type of empirical relation one wishes to uncover. In a cross-section of countries, if one is interested in the causal impact of participation, one can use standard 2SLS-3SLS estimation problem with the hope of finding appropriate instruments to determine labor market participation. Since variations in participation, both cross-country in time are too a large extent driven by female participation, the discussion that follows is based on finding instruments for female participation. See e.g. Wasmer (1997) for an early discussion of such instruments (divorce rate, the fraction of female in National Parliaments) and an illustration, and the recent works of Jaumotte (2003), Genre et al. (2004) and Pissarides et al. (2005) for far more detailed studies of the determinants of female participation. Notably, Jaumotte (2003) shows that marginal tax rates of second earner are one of the significant determinant of female activity rates. Other works indicates that the religion and church attendance tend to affect the aggregate share of women in the labor force (Algan and Cahuc, 2004) while Pissarides et al. (2005) find little effect of religion but an important role of startup costs.

If one is concerned by investigating causal links for a given country, one needs to deal with VAR and cointegration analysis and test Granger non-causality in using the lag structure. This is what we shall do now.

4.2 Methodology

Let p_t and u_t the two times series we consider and $\mathbf{X}_t = (p_t, u_t)'$. We estimate the following VAR system:

$$\mathbf{X}_t = \mathbf{D}(L)\mathbf{X}_{t-1} + \mathbf{Z}_t + \boldsymbol{\varepsilon}_t \tag{7}$$

where $\mathbf{D}(L)$ is a matrix of lags, $\mathbf{Z}_t$ is a set of additional controls (trends and constants) and $\boldsymbol{\varepsilon}_t = (\varepsilon_t^p, \varepsilon_t^u)'$ a vector of residuals. We then adopt the structural VAR approach and decompose $\boldsymbol{\varepsilon}_t$ into white noise structural innovation $\mathbf{u}_t$, i.e.

$$\boldsymbol{\varepsilon}_t = \mathbf{C}\boldsymbol{\xi}_t$$

where $\mathbf{C}$ is a 2x2 matrix. Result 2 above on the absence of a long-run impact of labor supply on unemployment could be used as an identification restriction, in the spirit of Blanchard and Quah (1989). They investigated the effect of technology and demand shocks in a bivariate VAR in assuming the absence of long-run impact of demand shocks on output and unemployment. We do not chose this road here, because, as argued in result 3, changes in the composition of the labor force, having affected the labor markets over more than three decades, can have a long-run impact given the relatively small sample (45 to 48 years) used. We thus restrict our analysis to short-run identification restrictions. The issue here is to disentangle structural shocks leading to a negative covariation between unemployment and participation (such as productivity shocks, demand shocks) and structural shocks leading to a negative covariation between unemployment and participation (such as shocks on the incentive to participate, such as the $\theta' s$).

Without loss of generality, we can write

$$\begin{aligned} \varepsilon_t^p &= C_{pp}\xi_t^p - C_{pu}\xi_t^u \\ \varepsilon_t^u &= C_{up}\xi_t^p + C_{uu}\xi_t^u \end{aligned} \tag{8}$$

where we expect all constants C_{ij} to be positive for $i = p, u$.

4.3 Structural identification

Given that the variance-covariance matrix $\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t'$ is obtained from the data, this imposes three restrictions on C_{ij}. We can thus only chose one coefficient, but off-diagonal terms can be identified from theoretical reasoning.

Indeed, the interpretation of C_{up} is how much unemployment will result from a structural innovation in participation. Having a prior on this coefficient is easier. In most countries, a fraction χ of the flows from non-participation to participation occur through unemployment, the complement $1 - \chi$ being direct flows into employment. In practice, χ is close to 0.5 in many countries. As argued in Garibaldi and Wasmer (2005), a part of direct flows from non-participation to employment (NE) is actually due to undetected infra-marginal transitions not detected by statisticians. An innovation ξ^p implies $\Delta L = \xi^p P_{15-64}$ additional participants, a fraction $1-\chi$ of them are directly employed. The NU flows (from non-participation to unemployment) is composed of the $\chi \Delta L$.[7] To compute the average increase in unemployment rate over a year, we use equation (6) with the quarterly value of α implied by the calibration. This implies that in that equation, $T = 4$. So, overall, taking for ΔU an average of the increase in unemployment at the time of the shock and one year after, we have that $\Delta U = \xi^p P_{15-64} \chi (\frac{1-e^{-\tilde{\alpha}T}}{\tilde{\alpha}T})$ so that

$$C_{up} = \frac{\chi}{p} \left(\frac{1-e^{-4\widetilde{\alpha}}}{4\widetilde{\alpha}} \right) \tag{9}$$

The interpretation of C_{pu} is the decline in participation when labor market conditions worsen, an effect known as the "flexion effect". We argue that for each increase in the number of unemployed worker ΔU due to a structural innovation in ξ_t^u, a number $\delta \Delta U$ of labor market participant return to inactivity. The parameter δ reflects an underlying labor supply curve for female workers, but where the determinant of participation rate is the rate of unemployment instead of the wage as in a standard labor supply curve. Thus, $\varepsilon^p = -\delta \Delta U / P_{15-64}$ where P_{15-64} is the 15-64 year old female population and $\Delta U = \xi^u (L + \xi^u) = \xi^u L$ ignoring second order terms where L is total active population. We have thus

$$C_{pu} = \delta p \tag{10}$$

There is however little knowledge about the dependence of participation to unemployment reflected in δ.

So, at this stage, one needs to make a choice about which restriction to impose, C_{up} or C_{pu}. Given that the labor supply parameter δ is unknown, we chose to impose a short-run restriction on C_{up} and recover C_{pu} from the data, which then allows us to obtain an estimate for the value of δ.

Note also that so far, we assumed that u referred to total unemployment rate and p referred to total participation rate. However, the methodology for the structural identification can be adapted to different specifications. Denote by L^f the number of female participants and by ω^f the share of women in the labor force, $L = L^f/\omega^f$. If we replace p by p^f in the VAR, we have to replace p by p^f/ω^f in equations (9)-(10), at least to the extent that no male job is displaced after an increase in female participation. If in addition, for each

[7] In GW (2005) we defended the view that a large part of observed NE flows is due to a statistical illusion, and decided to build a model where they should not be here. In this "purist interpretation", χ must be close to 1 instead. A positive value accounts for other interpretations of NE flows (I thank a referee for raising the point), such as the "jobs bump into people" mechanism also discussed in GW (2005). Being agnostic on that, I re-run the exercise with various values of χ ranging from 0.25 to 0.75 and found the same qualitative results, with quantitative changes, except for Italy, as discussed below.

Average 1956-2003	US	Fra	Ger	Ita
Unemployment rate u	5.73	6.18	4.48	8.20
Participation rate p	61.3	56.2	57.5	61.0
Female participation rate p^f	$47.7^{(a)}$	42.5	42.6	$40.1^{(b)}$
Share of female workers in labor force ω^f	39.7	39.4	39.1	$33.4^{(c)}$
Average 1995-2003				
Unemployment rate u	$4.85^{(d)}$	10.52	8.66	10.6
Participation rate p	65.3	55.3	57.1	60.6
Female participation rate p^f	58.2	48.4	48.2	45.9
Share of female workers in labor force ω^f	45.9	45.5	43.5	37.9
(a): 1956-2002; (b): 1959-2003 ; (c): 1958-2002 ; (d): 1995-2002				

Table 2: Sample statistics

additional female participant, there are Λ male workers displaced, we need to replace, in the same equations, p by $p^f(1+\Lambda)/\omega^f$. Following Blanchard and Perotti (2002) in a similar exercise, we compute C_{up} from the average value of p^f and ω^f over the sample, and, in unreported robustness checks, use instead the extreme values p^f/ω^f over three sub-periods without qualitative change.

5 Estimation and results

5.1 Data

We use time series data of the OECD Labor Force Statistics between 1956 and 2002. See Appendix B for a brief description. Data are yearly. Quarterly data are available since the early 90's, but as a first pass, we investigate only the longer-run data at the loss of higher frequency. Figure 3 displays the statistics for four countries. Table 2 shows the sample average of unemployment rates, participation rates and shares ω^f for the four countries of the study, in percent.

5.2 Specification and specification tests

Since we are in what follows interested in the estimation of some coefficients and short- and medium-run impulse response function, the question about whether the initial series are I(1) is not very important here (Sims et al. 1990). We decided to estimate equation (7) with a trend. A second specification choice involves the number of lags. We run the regression with one, two or three lags. It is only in the last case that all the autocorrelation of the residuals disappear. The larger the number of lags, the larger the number of parameters to estimate, so that the baseline specification has thus a trend and three lags.[8]

[8] I carried out formal tests such as the sequential modified likelihood ratio (LR) that the l^{th} lag is jointly zero. Depending on specifications and the criteria (Aike, Schwarz, Hanna-Quinn), the number of lags suggested was 2 or 3 (for most specifications) and in a few cases, 4. When this occured, this was not uniform result across criteria, and other criteria gave 2 or

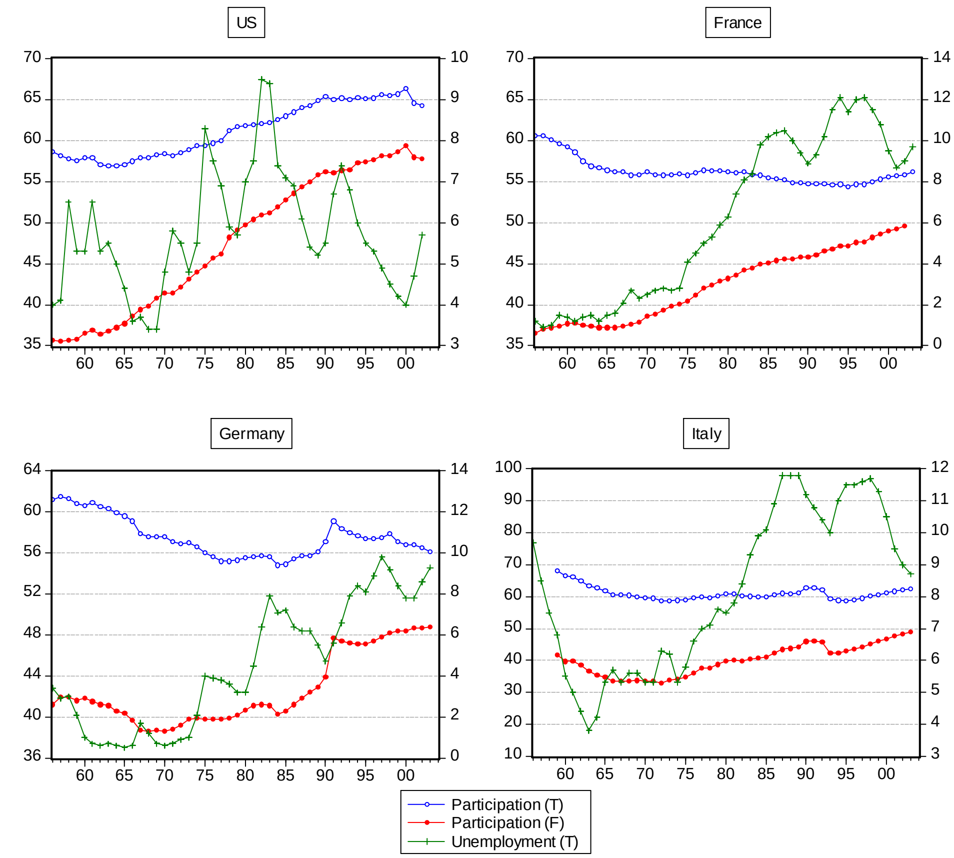


Figure 3: Total and female participation rates (left-axis, % of 15-64 population) and unemployment rates (right-axis, % of labor force).

The table in Appendix C reports some additional statistics of the VAR estimates. In general, the roots of the system are in the unity circle and usually below 0.90: the first line for each specification reports the highest modulus of the roots of the characteristic polynomial. The two figures in Appendix C also indicate the position within the circle of all roots. The second line report the covariance matrix of the estimated residuals. The third line tests for the presence of autocorrelation at the forth or fifth order. It appears that, for specification 1, residual autocorrelation subsists only for Italy. For specification II, there remains some autocorrelation at order 4 for the US, France and Italy.[9]

5.3 Identification of the shocks

Then, we proceed with the identification of the structural shocks and the impulse response based on C_{up}, based on the calculations of equation (9) and on sample averages in table (2). The value of χ is 0.5 for all countries and both specifications, except for Italy in specification I, for reasons made explicit below.

In the first specification $X = (p, u)'$, we use the sample average of p and the calibrated value of $\widetilde{\alpha}$ is 0.636 for the US-style economy and 0.311 for Continental European countries. In Table 3, the first line provides the country value of the short-run restriction C_{up} which ranges from 0.296 in the US to 0.510 in France. The second line uses the information on the variance-covariance matrix of $\boldsymbol{\varepsilon}$ to obtain the implied value of C_{pu} knowing C_{up} and the standard error. The third line uses equation (10) to obtain the implied value of δ and a standard error.

Given the structure of residuals of the VAR, this leads to the estimates of C_{pu} to the second line (s.e. in parentheses), between 0.206 for both France and the US to 0.498 for Italy. Italy is a specific case: with $\chi = 0.5$, the implied value of C_{up} would be 0.469. However, given the residuals of the VAR, it is impossible to invert the system (8) for values of C_{up} above 0.450 in Italy: the s.e. goes to infinity. A possible interpretation is that there are many NE flows in Italy due to inactivity being disguised shadow activity (Boeri and Garibaldi 2003). This should reduce χ and thus C_{up}. There is no way to have a better guess on χ because generally speaking, EN flows are noisy, subject to various statistical and theoretical interpretations as discussed in footnote 7. We take an agnostic view instead and propose to reduce χ to 0.4. To check the importance of this ad hoc assumption, the last column proposes an alternative specification with $\chi = 0.3$ with no qualitative implication on impulse response.

The implied value of δ (i.e., the number of participants leaving the labor force for each additional unemployed worker) is 0.34 in the US., a bit higher in France (0.37), larger in Germany and much larger in Italy (0.88). Note however

3. We thus set the number of lags to 3.

[9] For specification I, the inspection of the autocorrelogram shows that this is due to a large negative residual of the participation series in 1993. For specification II, a similar negative residual in 1993 is involved for Italy. For all countries, we checked that autocorrelations between leads and lags of residuals are always within two standard error bounds. This is indeed the case, except in one of the twelve checked lags for Italy, again due to year 1993. To remove this 4^{th} order autocorrelation, we would have to increase the number of lags to four, at some efficiency cost.

	US	Fra	Ger	Ita	Ita (2)
Specification I: $X = (p, u)'$		($\chi = 0.5$)		($\chi = 0.4$)	($\chi = 0.3$)
Restriction for C_{up}(new)	0.296	0.510	0.498	0.375	0.282
Estimated value of C_{pu}	0.206 (0.061)	0.206 (0.024)	0.383 (0.066)	0.535 (0.073)	0.451 (0.078)
Implied value of δ	0.34 (0.10)	0.37 (0.04)	0.67 (0.09)	0.88 (0.12)	0.74 (0.13)
Specification II: $X = (p^f, u)'$			($\chi = 0.5$)		
Restriction for C_{up}	0.226	0.398	0.394	0.358	
Estimated value of C_{pu}	0.166 (0.069)	0.117 (0.031)	0.330 (0.079)	0.565 (0.101)	
Implied value of δ	0.14 (0.06)	0.11 (0.03)	0.30 (0.07)	0.47 (0.08)	

Table 3: Identification restrictions

that given the standard error (0.09 in Germany and 0.12 in Italy), the difference between the two countries is not significant: δ could be below 0.76 in Italy with good probability. In fact, the high values of δ for these two countries is consistent with Figure 3 where total participation rates appear to be much more cyclical, and also consistent with the view that female workers in Germany and Italy have historically played, to a greater extent, a role of buffer to absorb labor demand shocks.[10]

This is what specification II aims at testing, in estimating a VAR with the vector $X = (p^f, u)'$. We have to adjust some parameters of the previous exercise. First, as argued above, we replace p by p^f/ω^f. Second, we need to have an estimate of how many workers (the parameter Λ) are displaced after an innovation in the number of female participants. In the absence of knowledge, we proceed as follows. We are going to estimate it from the US series, with the following additional identification assumption: we expect δ for the US in specification II to be about half of δ in specification I. Indeed, δ in specification II is the total number of female workers leaving the labor force after a change in the ξ^u — which is an increase in total unemployment. This implies that Λ is slightly above 0.5 but we round it to $\Lambda = 0.5$. We then assume that Λ takes the same value in the three other countries and recover the values of C_{pu} from the residuals and subsequently of δ. It appears that δ is relatively small for France but larger in Germany and Italy. As a robustness check, other regressions were run, adjusting the hiring rate α for women to lower values. Unreported results show that reducing α by 50% does not affect the value of δ by more than 2 or 3 percentage points and leave impulse response qualitatively unchanged.

[10]Note that $\chi = 0.3$ reduces further δ in Italy, which probably means that indeed χ must be lower than 0.5.

5.4 Impulse responses

Figure 4 displays the off-diagonal impulse response for a one standard deviation innovation and the confidence interval with two standard errors.[11] The first column of charts indicate that a positive unemployment shocks leads to a medium run decline in participation in the US (4 years) or Germany (6 years). The degree of persistence is larger in Italy and especially in France. The reverse causality, i.e. the response of unemployment to participation shocks, is in the second column of charts. It indicates that participation shocks appear to be rather persistent, except in Germany. The reason is that participation adjusts downwards rather fast, consistently with the estimates of δ above. Such is not the case in Italy, because of a higher persistence of the unemployment series. Finally, both France and the US exhibit significant persistence in unemployment after an innovation ξ^u, but the magnitude of the impact in France is twice as large after two or three years. Figure 5 represents the result of the same exercise for specification II, i.e. replacing p by p^f. Qualitatively, the results (variations and persistence) are the same as for specification I.

Interestingly, in both specifications, the confidence intervals always include zero in the long run. This allows for a further conclusion: Result 3 (long-run impact of heterogenous labor supply shocks) is not consistent with the data, while Result 2 (no-long run impact of homogenous labor supply shocks) is consistent with the data.

5.5 Variance decomposition

Before discussing cross-country differences, we carry another exercise, the variance decomposition for each series. This is reported for each specification in Tables 4 and 5. The variance decomposition for years after year 5 is not very informative, as most of the impact is no significantly different from 0, so we mostly focus the discussion on the first years after the shock. As shown in Table 4, in year 2 for instance, unemployment is mostly determined by participation shocks in Europe in the short run (70% for Germany, 50% for Italy, 93% in France), and much less (16%) in the US. This amounts to saying that total employment does not absorb the labor supply shocks in the short run in Europe, while in the US the opposite happened. A similar US-Europe divide appears for the participation series: unemployment shocks in year 2 determine the largest share of participation (98% in France, 83% in Germany and 95% in Italy), while it is only 24.5% in the US. Similar conclusions emerge when inspecting Table 5 for female participation. This suggests that participation changes in the US were driven by migration or demographic changes, while in Continental Europe, labor supply reacted more to economic conditions. As already stressed in introduction, the specification includes a country-specific trend, so that these results

[11] Inspection of on-diagonal terms, that is, the response of unemployment to innovations ξ^u or participation to innovations ξ^p indicate that shocks are not persistent in Germany, participation shocks are persistent in the US and in France, but there, zero is in the confidence band. In Italy, the confidence interval include 0 after 2 to 4 years in Italy. The on-diagonal terms are available upon request.

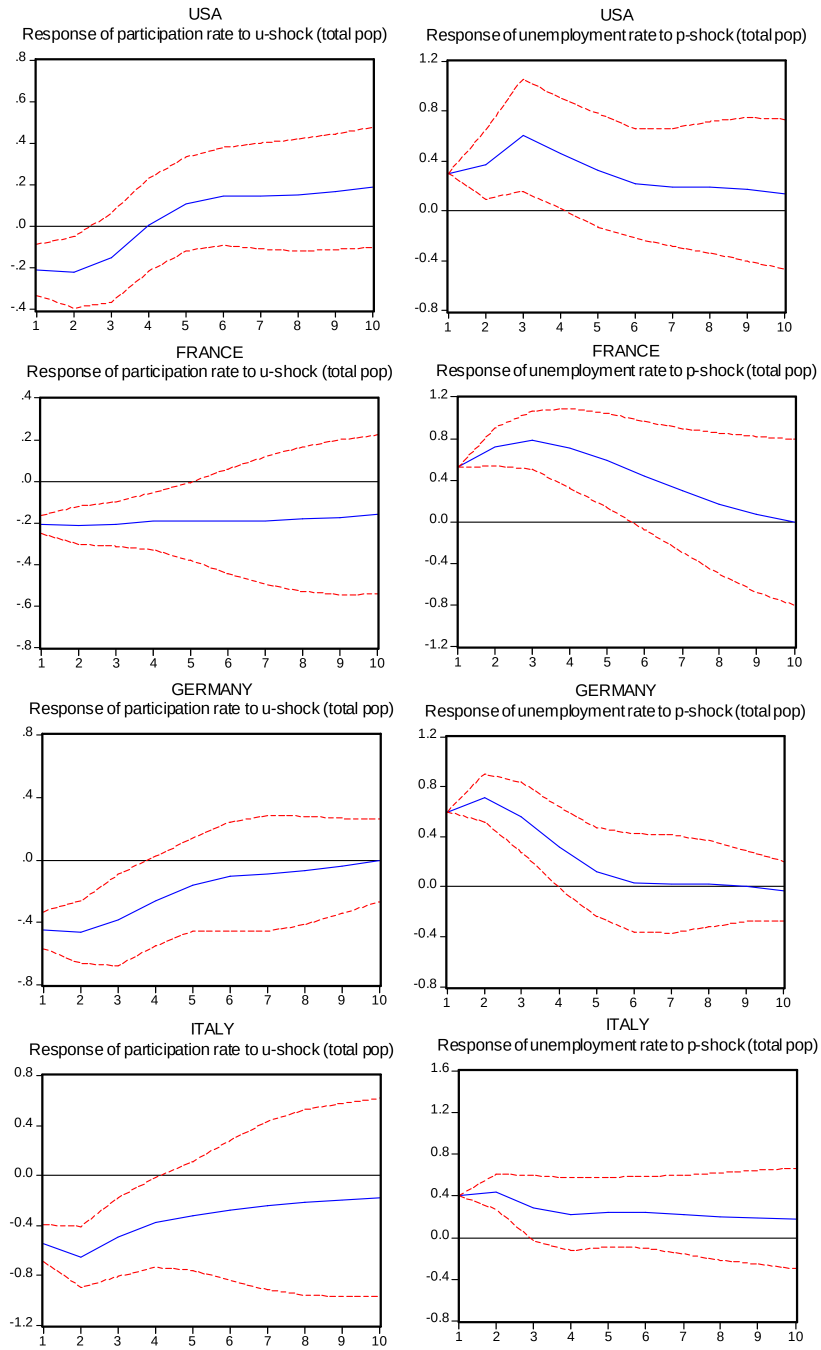


Figure 4: Response to one standard deviation structural innovation and +/- 2 standard errors confidence interval. Left charts: response of total participation rates to a structural innovation on total unemployment ξ^u. Right charts: response of total unemployment rates to a structural innovation on participation ξ^p.

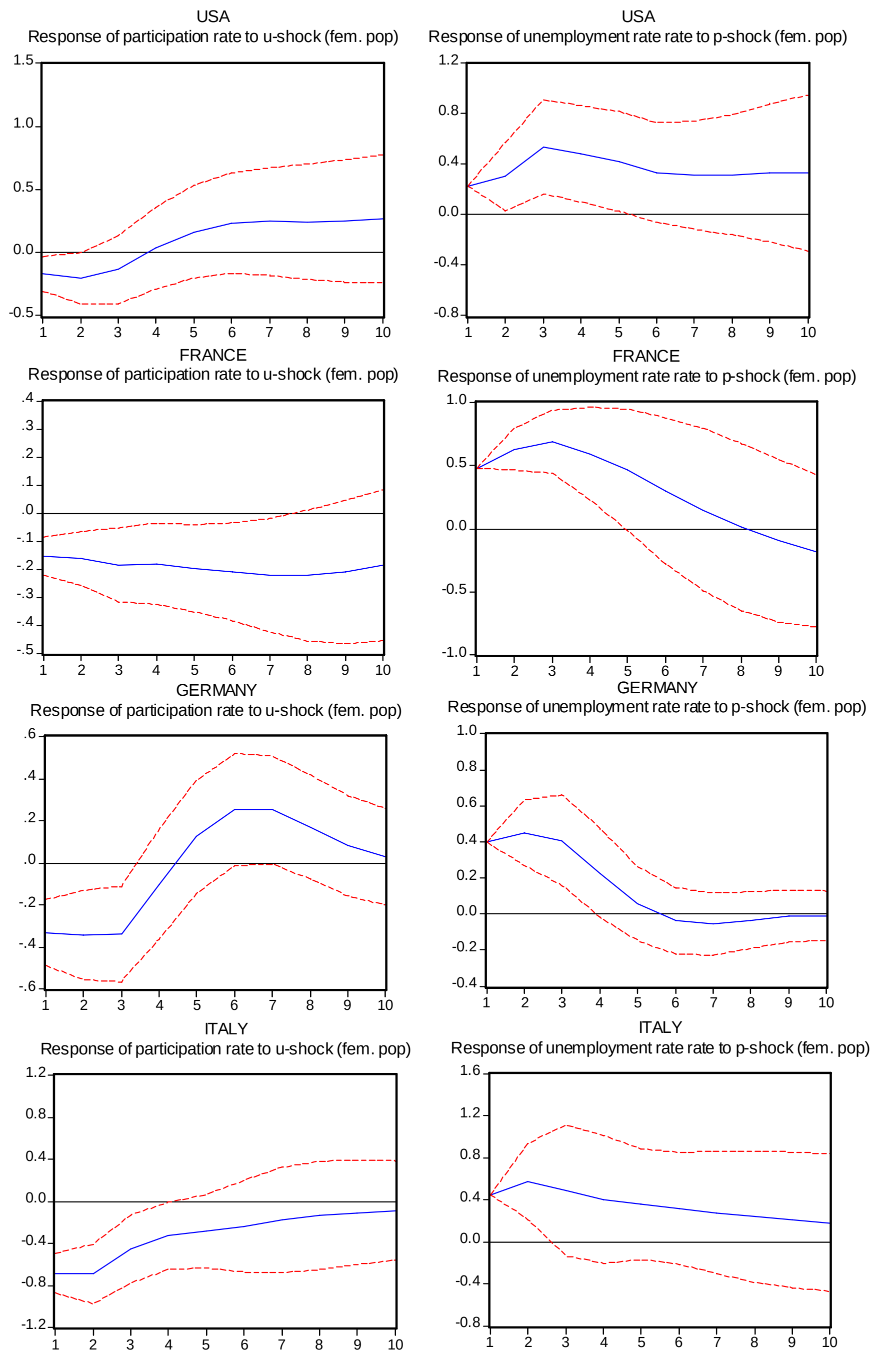


Figure 5: Response to one standard deviation structural innovation and +/- 2 standard errors confidence intervals. Left charts: response of female participation rates to a structural innovation on total unemployment ξ^u. Right charts: response of total unemployment rates to a structural innovation on female participation ξ^{p^f}.

	Usa		Fra		Ger		Ita	
Series	p		p		p		p	
Shock	ξ^p	ξ^u	ξ^p	ξ^u	ξ^p	ξ^u	ξ^p	ξ^u
Year 1	75.5	24.5	0.07	99.9	25.4	74.6	8.98	91.0
Year 2	70.9	29.1	2.27	97.7	17.0	83.0	4.83	95.2
Year 5	79.0	21.0	14.4	85.6	12.8	87.2	3.27	96.7
Year 10	80.2	19.8	52.7	47.3	28.5	71.5	3.11	96.9
Series	u		u		u		u	
Shock	ξ^p	ξ^u	ξ^p	ξ^u	ξ^p	ξ^u	ξ^p	ξ^u
Year 1	12.1	87.9	96.0	4.04	70.1	29.9	73.3	26.7
Year 2	16.3	83.7	93.0	7.00	67.8	32.2	50.5	49.5
Year 5	40.9	59.1	94.2	5.78	69.5	30.5	23.5	76.5
Year 10	40.0	60.0	94.6	5.45	63.3	36.7	19.8	80.2
Note: due to rounding errors, the sum of variances is not necessarily 100%								

Table 4: Variance decomposition, total participation and total unemployment

do not imply that participation is the major cause of unemployment, but simply that in the short-medium run, employment does not react to participation shocks in Europe.

6 Conclusion

This paper is concerned by causal links between unemployment rates and participation rates. A theoretical model allows to understand the causality from unemployment to participation decisions, and more precisely, from the job finding rate to the entry decision of workers. We then discuss the issue of opposite causality, and attempt to identify structural innovations of participation in the data and their impact on unemployment. Overall, results indicate a mild and significant positive impact of total participation and female participation on unemployment in the short and medium run, in a horizon of five to eight years depending on countries.

We also find differences across countries: in Continental Europe, participation react more to unemployment which itself determines a higher share of unemployment variations in the short-run. A possible explanation is European Malthusianism: first, there are social pressures to reduce labor supply when unemployment increases, and second, employment is stable and insensitive to pure labor supply shocks. In the US, shocks to labor supply are more important, due to immigration and demography: employment reacts fairly well to these shocks, and fluctuations of unemployment are driven by other shocks (demand or technology). Conversely, an increase in unemployment in the US is not followed by Malthusian pressures, and so little of the fluctuations of participation are drive by unemployment.

	Usa		Fra		Ger		Ita	
Series	p^f		p^f		p^f		p^f	
Shock	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u
Year 1	86.7	13.3	30.9	69.1	41.9	58.1	7.08	92.9
Year 2	83.9	16.1	32.8	67.2	31.5	68.5	6.19	93.8
Year 5	90.5	9.46	36.5	63.5	27.6	72.4	4.75	95.2
Year 10	87.8	12.2	27.3	72.7	37.1	62.9	4.94	95.1
Series	u		u		u		u	
Shock	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u	ξ^{p^f}	ξ^u
Year 1	7.25	92.7	84.5	15.2	58.6	41.4	88.9	11.1
Year 2	11.2	88.8	79.5	20.5	56.3	43.7	67.8	32.2
Year 5	40.6	59.4	85.4	14.6	61.5	38.5	45.9	54.1
Year 10	49.0	51.0	73.9	26.1	60.9	39.1	45.6	54.4
Note: due to rounding errors, the sum of variances is not necessarily 100%								

Table 5: Variance decomposition, female participation and total unemployment

Future research should aim at disentangling the differential impact of male and female participation shocks: an interesting extension would be to estimate a tri-variate VAR with (p^m, p^f, u): the interest is to allow for a larger range of short-run effects, like the differential response of male and female participation to unemployment innovations, or the response of female participation to innovation in male participation and reciprocally. The cost is to have to impose more identification restrictions and notably the impact of a shock on female participation to male unemployment. Preliminary results for the US however suggests that, even when one imposes the same short-run impact on unemployment of innovations in participation, female participation raise unemployment for a longer period while innovations in male participation exhibit no persistent impact on unemployment. Additional complexities arise for other countries, as the unit root of the characteristic polynomial seem closer to 1.

Another extension is to explore further the role of participation shocks in macroeconomics. So far, our unemployment shock is a combination of demand and technology. The debate launched by Galí (1999) might be enlighted by adding a new source of unemployment fluctuations. Although our results are suggestive that the results for the US might not be affected much by participation, the converse may hold in Continental Europe.

Appendix

A Derivation of the model

The proof follows GW (2005). Let $W(\theta), U(\theta)$ and $H(\theta)$ be the asset values of workers with home productivity θ. We have, assuming that only non-participants enjoy home production:

$$\begin{aligned}
(r+\mu)W &= w(\theta)(1-\tau)+\mu\int_{\theta^{\min}}^{\theta^{\max}}\max\left(W',H'\right)dF(\theta') && \text{(A1)}\\
&\quad +\delta[\max\left(U,H\right)-W],\\
(r+\mu)U &= b+\mu\int_{\theta^{\min}}^{\theta^{\max}}\max\left(U',H'\right)dF(\theta') && \text{(A2)}\\
&\quad +\alpha[\,\max\left(W,U\right)-U],\\
(r+\mu)H &= \theta+\mu\int_{\theta^{\min}}^{\theta^{\max}}\max\left(U',H'\right)dF(\theta'). && \text{(A3)}
\end{aligned}$$

We have: $W(\theta^q)=H(\theta^q)$ and $U(\theta^\nu)=H(\theta^\nu)$, and further, U is a constant of θ, as well as W when $\theta<\theta^\nu$. This leads to

$$\begin{aligned}
\theta^\nu &= \alpha(W-U)(\theta^\nu) && \text{(A4)}\\
\theta^q &= w(\theta^q)+\mu\int_{\theta^{\min}}^{\theta^q}[W'-\max(U',H')]\,dF(\theta') && \text{(A5)}
\end{aligned}$$

Let J be the value of a job for a firm in a free-entry world where the value of a job vacancy is zero in equilibrium. Then, if wages are bargained so as to conventionally share the total surplus $J+W'-\max(U',H')$in shares β and $1-\beta$, we obtain

$$\begin{aligned}
J &= (1-\beta)\frac{\theta^q-\theta^\nu}{r+\mu+\delta} && \text{(A6)}\\
w(\theta^q) &= (1-\beta)\theta^q+\beta y(1-\tau) && \text{(A7)}
\end{aligned}$$

and replacing the equations (A6) and (A7) into (A4) and (A5) we obtain the three main equations (2), (1) and (5) in the text.

Participation rate is calculated from stock-flows equation in a steady-state, and follow Garibaldi and Wasmer (2003):

$$\frac{1-p_{ILO}}{p_{ILO}}=\frac{1}{\mu F(\theta^\nu)}\left\{(\mu-\mu F(\theta^\nu)+\alpha)u-(1-u)\frac{\delta}{\delta+\mu}\left[\delta+q+\mu F(\theta^\nu)\right]\right\} \quad \text{(A8)}$$

The latter expression is a bit complex, but can be simplified if we introduce the notations $\varepsilon=\mu F(\theta^\nu)$ and $z=\mu\left[F(\theta^q)-F(\theta^\nu)\right]$. Indeed, notice that $\mu-\mu F(\theta^\nu)=q+z$ and $\delta+q+\mu F(\theta^\nu)=\delta+\mu-z$. We can rewrite the expression above in as the sum of two terms, the first one being q/ε and the second term being a more complicated expression.

$$\frac{1-p_{ILO}}{p_{ILO}}=\frac{q}{\varepsilon}+\frac{1}{\varepsilon}\left\{(z+\alpha)u-(1-u)(q+\delta(1-\frac{z}{\delta+\mu}))\right\}$$

We now add to GW (2005) the following trick. Define by p_l a more "liberal" notion of participation including the marginally attached workers who do not have a job. This quantity has a simple expression: indeed, realizing that $\partial p_l/\partial t = \varepsilon(1-p_l) + qp_l$, we obtain the steady-state level of p_l as

$$p_l = \frac{\varepsilon}{\varepsilon + q}, \tag{A9}$$

The difference between p_{ILO} and p_l is thus of the order of magnitude of $(\theta^q - \theta^\nu)$. It follows that, to the extent that θ^q and θ^ν are close to each others, the difference between p_l and p_{ILO} is small, as numerical exercises below also indicate. Further, both participation rates p_l and p_{ILO} converge to the neo-classical participation rate $p^* = F(y)$ when $\theta^q - \theta^\nu \to 0$. Now, since we know that p_l and p_{ILO} differ only by the number of marginally unattached workers which is proportional to z, we can anticipate that we can simplify. Indeed, noticing that $\alpha u = (q+\delta)(1-u)$, we have

$$\begin{aligned}\frac{1-p_{ILO}}{p_{ILO}} &= \frac{q}{\varepsilon} + \frac{1}{\varepsilon}\left\{zu + (1-u)\frac{\delta z}{\delta+\mu})\right\} \\ &= \frac{q}{\varepsilon} + \frac{z}{\varepsilon}\left\{\frac{\delta}{\delta+\mu} + \frac{\mu}{\delta+\mu}u\right\}\end{aligned}$$

and we thus obtain the equation (4) derived in the text. Note that in the calibration exercise reported in the text, the value of p_l for Europe and the US is respectively 0.603 and 0.728. Both are indeed very close from p_{ILO}.

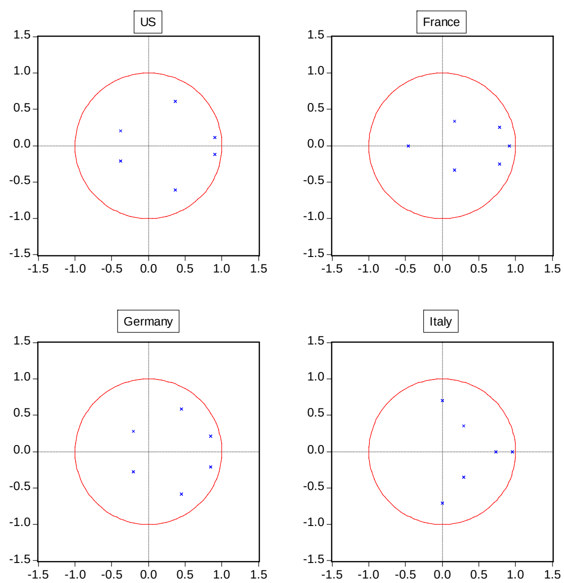
B Data description

Participation rates and unemployment rates are released by the OECD (Corporate Data Environment, Labor Force Statistics (part 3), http://www1.oecd.org/scripts/cde/default.asp, for all countries. Longest time series start in 1956 and end in 2003. Data for Germany include a structural break in 1991 at the time of reunification. Data for Italy have a structural break in 1993 at a time of a change in the design of the labor force survey and in the definition of unemployment. In addition, participation rates are not available between 1972 and 1976. It appears that total and female civilian population is indeed available, but the denominator of participation rates (the 15-64 population data) is not. The 15-64 Italian population was thus fitted for those years by interpolation and I reconstructed consistent time series of participation date. Regressions for Germany include a dummy variable D91=(year≥1991), and regressions data for Italy include two dummy variables, D93=(year≥1993) and D7276=(1972≥year≥1976).

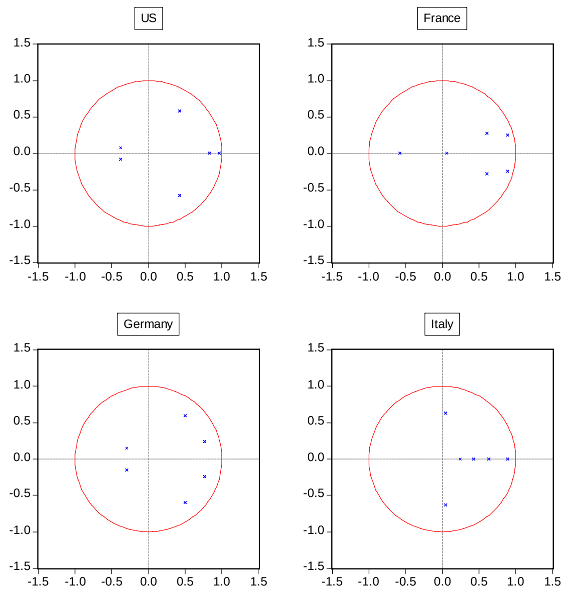
As regards to transition rates discussed in the calibration exercise, Auer and Cazes (2000) indicate, using LFS data that in France and Germany, the transition rate from employment to inactivity (*en*) over the period 1992-99 is roughly 4% a quarter, i.e. 1% a quarter, to which one should add regular retirement (an additional 0.5 to 0.7% a quarter), while the transition rate from employment to unemployment (*eu*) is about 1% a quarter . In Italy, the same authors find a similar *en* rate, 1% a quarter, and a much lower *eu* rate, about 0.5% a quarter. Nonetheless, Boeri and Garibaldi (2002), extend the traditional flows analysis to the shadow sector: they show in their Table 4 that flows from employment to shadow unemployment are as high as regular *eu* flows in the South, and about 40% of that flow in the North. Overall, the actual *en* flow rate in the three Continental European countries is 1%, to which we add 0.6% of retirement. The model tends to overshoot that rate *en*, while $\delta = eu$ is adequately chosen here. For the US, GW (2005) report monthly flows for the 25-54 population over the period 1995-2000. Translated into quarters, this amounts to 2.0% for *eu* flows and 2.46 for *en* flows to which some 0.5% retirement rate should be added, i.e. 3% a quarter. The calibration of the model is thus inaccurate in this dimension, a finding already discussed in GW (2005). An improvement of the calibration for the US would imply changing λ or addressing the infra-year transition rate.

C Specification tests

VAR specification tests				
	US	Fra	Ger	Ita
Specification I: $X = (p, u)'$				
Modulus of highest root of charact. polynomial	0.915	0.911	0.869	0.954
Covariance matrix $\widehat{\varepsilon}\widehat{\varepsilon}'$	$\begin{matrix} 0.18 & -0.05 \\ -0.05 & 0.75 \end{matrix}$	$\begin{matrix} 0.04 & -0.02 \\ -0.02 & 0.28 \end{matrix}$	$\begin{matrix} 0.20 & -0.01 \\ -0.01 & 0.37 \end{matrix}$	$\begin{matrix} 0.32 & -0.02 \\ -0.06 & 0.20 \end{matrix}$
Portemanteau test of residual corr. at lag 4/5[(a)]	8.57 / 9.51	7.45 / 11.9	4.82 / 5.59	13.3* / 22.2*
Specification II: $X = (p^f, u)'$				
Restriction for C_{up}	0.964	0.922	0.798	0.888
Covariance matrix $\widehat{\varepsilon}\widehat{\varepsilon}'$	$\begin{matrix} 0.21 & -0.04 \\ -0.04 & 0.73 \end{matrix}$	$\begin{matrix} 0.03 & 0.02 \\ 0.02 & 0.27 \end{matrix}$	$\begin{matrix} 0.27 & 2.10^{-3} \\ 2.10^{-3} & 0.38 \end{matrix}$	$\begin{matrix} 0.47 & -0.02 \\ -0.02 & 0.47 \end{matrix}$
Portemanteau test of residual corr. at lag 4/5[(a)]	12.4*/13.3	10.8*/18.2*	9.25/11.1	13.3* / 23.3*
(a): Q-value ($\chi^2(4)$ and $\chi^2 2(8)$ resp.) ; * at the 5% level.				



Appendix. Position of the inverse roots of the AR characteristic polynomial, specification I.



Appendix. Position of the inverse roots of the AR characteristic polynomial, specification II.

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