UNEMPLOYMENT AND FERTILITY: A LONG RUN RELATIONSHIP

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Abstract

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Keywords: Unemployment rate; fertility rate; long memory; fractional integration; fractional cointegration

JEL Classification: C22; C32; J13; J64

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1. Introduction

In the last 100 years, with the emergence of women in the labor market, it is necessary to study the gap between the intention of having children and actually having them. The persistent gap between desired and achieved fertility has stirred concerns about unhappy citizens underachieving their fertility goals, and it has provided a strong argument in favor of social policies aimed at removing obstacles such as unstable working conditions or difficulties in combining family and work (Testa, 2014). This incorporation of women into jobs across Europe has led to a decline in fertility (Adserá, 2018). Although the fall is widespread throughout Europe, not all countries reflect the same incidence.

In countries such as Greece or Spain, and in general in southern Europe, fertility is lower than in the rest, mainly because states do not help future parents and jobs are more precarious and have worse conditions. Other countries, such as Austria, Germany or Switzerland do not have a model of family-work balance, which delays the decision to have children, as parents would need to keep working in order to maintain the child's needs (Hoen and Andersson, 2017). In Belgium, however, we observe a perfect model of family life and work balance, a fact that immediately encourages increased fertility. Women in countries where more women are employed have been accommodated with more extensive family service provisions. This would explain why fertility rates are higher in Western Europe where family services are more extensively funded and available than in Eastern Europe where family services are scarce. For Eastern European countries, we also identified countervailing forces that might influence the fertility gap: poor economic situations and difficulty combining work and family (both being conducive to small family size (Beaujouan and Berghammer, 2017). In general, we can say that those women with more studies prefer to devote more time to a single child than to having several and that leads to a decrease in birth rate (Testa and Stephany, 2017).

This is a consequence of society continuing to require women to take care of their children, thereby limiting the horizons of the mothers, even more so if they have a good job and a good salary, so they decide to delay having a family (Bellani and Esping-Andersen, 2013) when it should be just the opposite, a good position and good salary affording a woman the freedom to choose to start a family. With regard to men, the effect is the other way round: the higher the salary is and the better working conditions are, the more the decision to start a family is encouraged and therefore the fertility rates increase.

If we focus on how unemployment affects men and women, there are different opinions. Some think that improving female employment lowers fertility (Adserá et al., 2012) and others believe that only the role played by men is really decisive when it comes to increasing or decreasing fertility (Kohler, Billari and Ortega, 2002). It is true that not having a stable job or being unemployed causes economic instability that does not invite you to start a family or have more children (Sobotka et al., 2011). Nor do precarious and part-time jobs help. In Eastern European countries when deciding whether to have children the type of employment and the general stability of the family economy is taken into account. In contrast, in other more advanced countries, once the woman enjoys more privileges and higher salaries, it is she who ultimately makes this decision (Cuestas et al., 2015).

The case of the US and Japan are curious in that it is observed that not the absence of social security or having poorly paid employment or jobs with little chance of getting salary increases or promotions, cause fertility to decline (Esping-Anderson, 2017; Juhn and McCue, 2017). In conclusion, not having a permanent job or having a precarious job is decisive when having a family (Schmitt, 2008).

Other factors such as the division of domestic work within the family (McDonald, 2006), and the education received (Rindfuss and Brewster, 1996) should also be considered. It has been observed as a general rule that when women have greater access to higher and university studies, by increasing their education, they ultimately achieve higher salaries and job responsibilities and therefore greater autonomy for decision-making (Kreyenfeld, 2010) and this implies a lower fertility rate. This incorporation of women in the labor market means that the number of marriages is falling and that the coexistence of existing ones is becoming more difficult, and the number of divorces are increasing (Goldstein et al., 2013). Fortunately, starting in the 90s, the role of the working woman began to be valued and with this the fertility rates improved slightly (Thévenon, 2010).

Once it is accepted that women work inside and outside the home, we face another decisive factor when assessing fertility rates: the division of domestic work. Those women who gradually reach positions of greater responsibility in their work, delay their decision to have a child (Goldin, 2014). If women duplicate their workload, working outside and inside the house, they will seriously consider whether to fully commit to their professional career or to start a family (Autor et al., 2015).

Emancipated women, greater labor promotion, divorce, abortion and contraceptives have allowed women to delay the decision to have children (Coleman et al., 2002). Fertility rates, according to other studies, are also influenced by a decisive factor in continuing socio-economic development: gender equality (Myrskylä et al., 2009).

This paper investigates how fertility and unemployment are related, testing if one of the variables has an influence over the other in the long run. These variables are key to identifying the extent to which labour market may help reduce (or increase) the potential opportunity cost of a new birth. To test the relationship between these two variables we use techniques based on concepts such as fractional integration and cointegration in

Europe, Japan and the US. These countries are characterized by very low fertility and a prolonged economic downturn during which both unemployment and non-standard employment increased markedly among young adults.

2. Literature review

Studies of the low-fertility phenomenon very much emphasize the influence of women's job insecurity, unemployment risks, and difficulties of reconciling work and motherhood (Kohler et.al., 2002; Adsera, 2004; Kreyenfeld, 2010). High rates of unemployment and difficulties in attaining a stable job (Adsera, 2004), are some of the factors hypothesized to drive the so-called «syndrome of delay» (Livi Bacci, 2001). This is especially the case for transitions like parenthood that require long-term binding commitments (Esping-Andersen, 2013).

The increase in women's labour force participation has been accompanied by a steady decline in fertility. Yet again, we find important reversals both at the macro and micro levels. Ahn and Mira's (2002) study shows that at the macro level, the traditionally negative relationship between female labour force participation and fertility rates has turned positive since the mid-1980s. The trend is just the opposite regarding women's unemployment: the cross-country correlation shifts from positive to negative (Esping-Andersen, et al. 2013). At the individual level, however, the association between female labour force participation and fertility tends to be negative, although there are important variations across cohorts and across countries (Matysiak and Vignoli, 2008). The impact of women's employment on childbearing is positive in Northern Europe (Andersson, 2000), but negative in Southern European countries (Baizán, 2005).

There are different key obstacles to fulfilling fertility preferences (Fernandez-Crehuet et al., 2017). Firstly, macro-level conditions related to labour market structures and opportunities matter. Since stable employment has become a pre-requisite for childbearing, the high unemployment rate of young adults and the unstable position of many of those employed are clearly major obstacles to childbearing. Secondly, the institutional and policy setting also matters. Public support for women and men to combine paid work and family responsibilities has never been a priority in some countries (Fernandez-Crehuet et el., 2016). Most policies have not gone beyond abstract commitments, strong rhetoric and piecemeal interventions. The current economic crisis, with rising unemployment and job insecurity, and the implementation of austerity programs make it even more difficult to envision more comprehensive support for families in the near future (Castro-Martín and Martín-García, 2013).

There are also multiple approaches when modelling these two variables (i.e., unemployment rate and fertility rate) including those based on cross-section, time series and panel data. Focusing on time series, the literature on modelling unemployment is very extensive and most authors agree that unemployment is a highly persistent variable, supporting the hypothesis of hysteresis at least in the European case (Blanchard and Summers, 1986; Chang 2011; García-Cintado et al., 2015; Munir and Ching, 2015; Cuestas et al., 2015; Caporale and Gil-Alana, 2018; etc.).

There are a large variety of models used when describing unemployment rates. Among them, the AutoRegresssive Moving Average (ARMA) and all its variants (ARIMA, SARIMA, ARFIMA, etc.) are the most popular ones based on their simplicity and good forecasting abilities. Other standard models are the Bayesian Structural Time Series (BSTS), the Smooth Transition AR (STAR) and other more complex non-linear approaches.

Fractional integration analysis provides us with greater analytical flexibility: by estimating the value of the differencing parameter, studies can make an assessment about the validity of alternative theories of unemployment. Recent contributions, Gil-Alana (2001a,b, 2002) and Caporale and Gil-Alana (2007, 2008), among others, conclude through applying ARFIMA models, that the structuralist view is more appropriate as a characterization of European unemployment.

Caporale and Gil-Alana (2007) proposed a model of the US unemployment rate which can account for both its asymmetry and its long memory. Their approach, based on the test of Robinson (1994), introduces fractional integration and non-linearities simultaneously into the same framework, unlike earlier studies which employ a sequential procedure (see van Dijk and Franses, 1999). They found out that the order of integration of the series is higher than 1, implying that, even when taking first differences, the series still possess a component of long memory behavior.

Another group of papers analyze the order of integration of unemployment rate by means of unit root tests for panel data, in order to consider cross-sectional information. Thus, Song and Wu (1997, 1998) and León-Ledesma and Thirlwall (2002) find that the hysteresis hypothesis is supported by EU data, whereas the NAIRU theory is more appropriate to characterize US unemployment. On the other hand, Christopoulos and León-Ledesma (2007) find evidence against the hysteresis hypothesis for EU data. However, the issue of structural breaks is not considered by these authors. Other authors who do apply panel unit root tests with structural breaks (Murray and Papell, 2000, and Strazicich et al. 2001), find more evidence supporting the structuralist theory of unemployment.

Caner and Hansen (2001) studied a TAR model of two regimes with an autoregressive unitary root developing a theory that permits us to distinguish between non-linear and non-stationary processes. Their study concludes that the unemployment rate in the US is a stationary nonlinear autoregressive threshold. Skalin and Tëravirta

(2002) used the L-STAR (Smooth Transition Autoregressive) model, arguing that the observed asymmetry can be captured by this simple model by introducing local non-stationarity into the calculation and thereby achieving more stable results. They establish that if the unemployment rate is a non-linear stationary process, the linear VAR (Autoregressive vector type model) will be erroneous since they assume that it is an I(1) variable and includes cointegration relationships along with other variables.

Focusing on fertility, early works relate female income effect on fertility in the sense that women choose to have fewer children as a consequence of an increase in the economic costs of childrearing (Easterlin, 1973, Mincer, 1963). In fact, the relationship between fertility and economic conditions has been widely examined (see, e.g., Butz and Ward, 1979; Macunovich and Easterlin, 1988; Rindfuss et al., 1988; etc.) and, unemployment has been the most commonly used measure for this economic conditions (Kreynfeld, 2010; Adsera, 2011; Goldstein et al., 2013; Amialchuk, 2013; etc.). Dealing with the relationship between the two variables, a number of authors argue that the fall in fertility rates is associated to high level of female unemployment (Brewster and Rindfuss, 2000; Esping-Andersen, 2009; Engelhardt and Prskawetz, 2004; etc.).

In a multivariate context and from a methodological viewpoint, the notion of cointegration arose out of the concern about spurious or nonsense regressions in time series. The problem is to find a way to work with two or more possibly nonstationary series in a fashion that allows us to capture both short run and long run effects. In more technical parlance, cointegration is the link between integrated processes and steady state equilibrium (Rajbhandai, 2016).

Another group of papers analyze the fertility rate in Europe. Some studies applied fixed-effect modeling (Allison, 2009). This model aims to identify causal mechanisms by exploiting within-country variations. Goldstein et al. (2013) used data from the Human

Fertility Database, Eurostat and OECD database to study how changes in unemployment rates affect birth rates across Europe. The dependent variable in their investigation was the age-specific fertility rate and they inserted a linear time trend to consider underlying fertility trend associated with postponement. They concluded that unemployment rates are closely associated with fertility rate development.

A variety of mathematical models have been proposed in order to describe the age-specific fertility pattern. Using data of the United Kingdom, Ireland and the US, Peristera and Kotaski (2007) proposed a parametric model in order to describe the fertility rate in these countries. In order to evaluate the adequacy of the model proposed, they fit the three alternative formulae to a variety of periods and cohort datasets of several populations. Furthermore, they compare these with other models already existing in the literature.

Cazzola et al. (2006) examined the relationship between fertility and unemployment in Italy. They used a monitoring approach for the identification of structural breaks in both time series and used a dynamic regression to identify specific temporal links between unemployment and fertility. They concluded that in some parts of Italy, the recent rise of unemployment is negatively correlated with the fertility rate. This paper is the closest one we find to ours since we also look at the relationship between unemployment rate and fertility rate. However, we use a very different time series approach based on concepts such as fractional integration and cointegration that are briefly presented in the following section.

3. Methodology

The methodology used in this work employs fractional integration and cointegration techniques. These methods belong to a broader concept named long memory or long-range dependence that means that observations are highly dependent across time and this dependence holds even between observations which are far distant in time.

We start this section by providing some definitions. Given a covariance or second order stationary process $\{u_t, t = 0, \pm 1, \ldots\}$ with autocovariance function, γ_k, and defined as:

$$\gamma_k \;=\; E\left[\left(u_t - \mu\right)\left(u_{t+k} - \mu\right)\right],$$

where $\mu = E(u_t)$ for all t, we say that u_t is short memory (or integrated of order 0) and denoted as I(0) if the infinite sum of all its autocovariances is finite, that is,

$$\sum_{k=-\infty}^{\infty} |\gamma_k| \;<\; \infty.$$

Within this category of short memory processes, we include the classical stationary AutoRegressive Moving Average ARMA-type of models.

On the other hand, we say that a process displays the property of long memory if the infinite sum of its autocovariances (or pseudo-autocovariances in case of nonstationary series) is infinite, i.e.,

$$\sum_{k=-\infty}^{\infty} |\gamma_k| \;=\; \infty,$$

and here we can consider the standard unit roots or I(1) processes widely employed in the literature. However, any I(d) processes with positive d satisfies the property of long memory and thus, fractional integration (i.e., when d is a fractional positive value) also belongs to this category.

We say that a process $\{x_t, t = 0, \pm 1, \ldots\}$ is integrated of order d, and denoted as I(d) if after removing its d-difference, the remaining process is I(0). In other words, x_t is I(d) if:

$$(1 - L)^d x_t \;=\; u_t\,, \quad t \;=\; 0, \pm 1, \ldots, \tag{1}$$

with $x_t = u_t = 0$ for $t \leq 0$, where L indicates the lag operator, i.e., $L^k x_t = x_{t-k}$, and with I(0) u_t. In this context, if u_t is, for example, an ARMA(p,q) process, x_t is said to be a fractionally integrated ARMA, ARFIMA(p, d, q) process.

As earlier mentioned, the classical cases examined in the literature impose d = 0 (stationarity) or d = 1 (nonstationarity), but the I(d) processes allow for a greater flexibility by permitting the differencing parameter d to be a fractional value. In such a case, the polynomial on the left-hand side of equation (1) can be expressed in terms of its Binomial expansion such that, for any real value d,

$$(1 - L)^d \; = \; \sum_{j=0}^{\infty} \psi_j L^j \; = \; \sum_{j=0}^{\infty} \binom{d}{j} (-1)^j L^j \; = \; 1 - \; d \; L \; + \frac{d\,(d-1)}{2} L^2 \; - \; \dots \; ,$$

implying that the equation in (1) can be expressed as

$$x_t \; = \; d\, x_{t-1} \; - \; \frac{d\,(d-1)}{2} x_{t-2} \; + \; \dots \; + \; u_t \, .$$

Thus, if d is a fractional value, x_t depends on all its past history and the higher the value of d is, the higher the dependence between the observations is, so d can be taken as a measure of the degree of persistence in the data. If d belongs to the interval (0, 0.5) x_t is still covariance or second order stationary though displaying long memory and with the shocks disappearing in the long run; if d is in [0.5, 1), x_t is no longer covariance stationary though it is still mean reverting, with shocks still having a transitory nature but presenting long lasting effects, while $d \geq 1$ implies lack of mean reverting behaviour.

Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981) proposed these models in the 80s and they have been widely employed in the modelling of time series since the late 90s starting with the paper by Gil-Alana and Robinson (1997). These authors examined an updated version of fourteen US macroeconomic series employed earlier by Nelson and Plosser (1982). Using ADF (Dickey and Fuller, 1979) tests, these authors found that the series were I(1), while Gil-Alana and Robinson (1997) extended the analysis to the fractional case, showing that the series were in fact I(d) with d being a fractional value, and statistically different from 1.

In the empirical section carried out in Section 5 we start with the univariate analysis investigating the order of integration of the two variables, the fertility rate and the unemployment rate. To allow the incorporation of deterministic terms and following standard approaches in nonstationary contexts (Bharghava, 1986; Schmidt and Phillips, 1992; etc.), we consider the following model,

$$y_t \;=\; \beta_0 \;+\; \beta_1 t + x_t; \qquad (1 - L)^d x_t \;=\; u_t\,, \qquad t \;=\; 0,\, 1, ..., \tag{2}$$

where y_t is each of the two observed time series; β_0 and β_1 are unknown coefficients referring respectively to an intercept and a linear time trend, and x_t is I(d) so that u_t is I(0) expressed in terms of a white noise process. We estimate and test the differencing parameter d throughout the Whittle function expressed in the frequency domain (Dahlhaus, 1989). We will employ a version of the tests of Robinson (1994) which is very convenient not only for the univariate analysis of the series but also when testing for long run relations among the two variables.

In the second part, we look at a long run equilibrium relationship between the two variables by using fractional cointegration methods, which is the natural generalization of the concept of cointegration (Engle and Granger, 1987) to the fractional case (Gil-Alana, 2003). This methodology consists of the following two steps:

a) testing the order of integration of the individual series (in our case, fertility rate, x_{1t}, and unemployment rate, x_{2t}), by using ADF (Dickey and Fuller,1979) tests, and

b) if the two individual series are I(1), testing the order of integration of the residuals from the cointegrating regression:

$$x_{1t} \;=\; \alpha \;+\; \beta x_{2t} \;+ u_t, \qquad t \;=\; 1, 2, \;... \;, \tag{3}$$

once more carrying out ADF tests with appropriately obtained critical values. Extending this approach to the fractional case, by assuming for instance that d is the order of integration of the individual series, and d-b the one of the potential cointegration regression, Cheung and Lai (1993) and Gil-Alana (2003) computed finite sample critical values for testing the null hypothesis of no cointegration (b = 0) against the alternative of fractional cointegration (b > 0).

4. Data

The data used in the paper consist of annual data of the unemployment rates and the fertility rates for Denmark, France, Ireland, Italy, Japan, Luxembourg, Netherlands, Portugal, Spain, Sweden, the UK and the US, for the time period from 1983 until 2017. We choose these countries because they contain the longest available datasets. The unemployment and fertility rates data come from Eurostat, the Statistical Office of the European Union, responsible for publishing high-quality statistics and indicators at the European level. Plots of the time series are displayed across Figures 1 and 2.

[FIGURES 1 AND 2 AND TABLE 1 ABOUT HERE]

As we can see from Table 1, the highest fertility rates correspond to Sweden (6.24), which is also the country with the lowest unemployment rates (1.81). Sweden is followed at a considerable distance by Ireland (2.03) and the US (1.96). We observe that some of the countries with low fertility rates like Spain (1.26) and Ireland (2.03), correspond to the highest unemployment rates, Spain (16.84) and Ireland (10.97). Seeing these values, we can imagine that the two variables are somehow correlated. In the next section we study how one variable affects the other, this being the hypothesis that we will be tested in the following section.

5. Empirical results

Table 2 displays the estimated values of d and the 95% confidence bands of the non-rejection values of d using Robinson's (1994) tests, in the model given by equation (2) under the three standard cases of i) with no deterministic terms (i.e., $\beta_0 = \beta_1 = 0$ in equation (2)), ii) with an intercept ($\beta_1 = 0$), and iii) with an intercept and a linear time trend (β_0 and β_1 estimated from the data). Then, we select the appropriate model (marked in bold in the table) by looking at the significance of the coefficients throughout their corresponding t-values in the d-differenced regression.[2]

The upper part of the table reports the results for the fertility rate while the lower one focusses on unemployment. The first thing we observe in the table is that the time trend is only required in the fertility rate for Ireland and Portugal, an intercept being sufficient in the remaining countries and also in all cases for the unemployment rates. If we focus now on the estimated values of d, and starting with the fertility rates, we observe that all values are above 1 except Luxembourg (0.97), though in this case, the unit root null cannot be rejected. The rest of the values range between 1.06 (Denmark) and 1.66 (Sweden), and while the unit root null (i.e., d = 1) cannot be rejected for Luxembourg, Denmark, Netherlands and Portugal, it is rejected in favour of d > 1 in the remaining countries.

TABLE 2 ABOUT HERE

Looking at the univariate results for unemployment (in the lower part of Table 2) we see that the values of d are now all above 1, ranging between 1.13 (France) and 1.86 (Netherlands). The unit root null cannot be rejected in the following cases: France (1.13),

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[2] Note that equation (2) can be expressed in terms of a single equation as $\tilde{y}_t = \beta_0 \tilde{1}_t + \beta_1 \tilde{t}_t + u_t$; where $\tilde{y}_t = (1-L)^{d_o} y_t$; $\tilde{1}_t = (1-L)^{d_o} 1$, and $\tilde{t}_t = (1-L)^{d_o} t$, and since u_t is I(0) by construction, standard t-tests can be applied for β_0 and β_1 in (2).

Italy (1.16), Luxembourg (1.23), Denmark (1.29) and US (1.37), and evidence of orders of integration above 1 is obtained for Japan (1.36), Sweden (1.49), UK (1.57), Portugal (1.64), Ireland (1.70), Spain (1.71) and the Netherlands (1.86). Thus, a conclusion that can be drawn from this table is that there is no evidence of mean reversion (or transitory shocks) in any of the series examined, since all orders of integration are found to be equal to or higher than 1 in the two variables for all countries under investigation.

Next, we look at the bivariate case, by looking first at the possibility of cointegration between the two variables of interest. Here, a necessary condition is that the two series statistically must display the same degree of integration. Table 3 displays the confidence intervals for the orders of integration of each country in the two variables.

TABLES 3 AND 4 ABOUT HERE

We see in Table 3 that the confidence bands overlap one each other for each country suggesting the equality in the degrees of integration. Nevertheless, we also conducted formal tests for this hypothesis, in particular, using Robinson and Yajima (2002), and identical conclusions were obtained with the method described in Hualde (2013). The results support the hypothesis of equal degrees of integration in all countries.

We next conduct Engle and Granger's (1987) approach, testing for cointegration between the two variables in all countries in the sample. Using this approach, the results are reported in Table 4. Based on the equality in the orders of integration of the individual series, we display in the table the estimates of d and the 95% confidence intervals on the estimated errors in the regression of one of the variables against the other for each country in the sample. Panel i) in Table 4 refers to the case of a regression of the fertility rate against the unemployment rate (i.e., $FR_t = \gamma_0 + \gamma_1 UR_t + x_t; \;\; (1 - L)^d x_t = u_t$, where FR refers to the fertility rate, and UR to the unemployment rate), while the reverse case, i.e., unemployment versus fertility ($UR_t = \gamma_0 + \gamma_1 FR_t + x_t; \;\; (1 - L)^d x_t = u_t$), is displayed in panel ii). The most noticeable feature observed in this table is that there is no a single case where mean reversion takes place since all the confidence intervals for d in the two panels include the value of 1. Thus, though we observe several cases with significantly negative slope coefficients γ_1 (such as France, the Netherlands, Portugal and the USA in panel i), and France, Italy, Portugal and the USA in panel ii)), this relationship can be taken as spurious based on the nonstationarity of the two variables (Hendry and Juselius, 2000; Gil-Alana and Solarin, 2018; etc.)

Our final approach consists of assuming that unemployment is weakly exogenous in relation with the fertility rate. In particular, we consider now the following regression model,

$$\mathrm{FR}_t = \delta_0 + \delta_1 UR_{t-1} + x_t; \qquad (1-L)^d x_t = u_t, \qquad t = 0, 1, ..., \qquad (4)$$

and jointly estimate δ_0 and δ_1 along with d in equation (4). Using this specification, we can still employ Robinson’s (1994) test for the estimation of the differencing parameter d with consistent estimates for the constant and the slope coefficients in equation (4).[3] The results using this approach are reported in Table 5. Panel i) assumes that u_t in (4) is a white noise process, while in panel ii) we allow for autocorrelation by using the non-parametric exponential approach of Bloomfield (1973) that approximate AR structures, and that accommodates very well in the context of fractional integration (see, Gil-Alana, 2004).

TABLE 5 ABOUT HERE

Starting with the case of white noise errors (panel i), Table 5) we observe that the unit root null hypothesis cannot be rejected in the cases of Denmark, Japan, Luxembourg, the Netherlands and Portugal, this hypothesis being rejected in favour of higher orders of

[3] Note that under this specification we are not testing for cointegration since UR_{t-1} is taken as weakly exogenous in the regression model (4).

integration in the remaining countries. Thus, we observe high degrees of persistence in the data. More importantly, the coefficients relating the two variables are found to be statistically significant in only four countries (the Netherlands, Portugal, Spain and the US), all them being significantly negative. Allowing for autocorrelation throughout the model of Bloomfield (1973) (panel ii), the results are qualitatively similar, though the confidence intervals of dare much wider; however, the coefficients relating the two variables is significantly negative in exactly the same four countries as in panel i), i.e., the Netherlands, Portugal, Spain and the US. Netherlands and US have a high share of women in part-time jobs. The availability of part-time work facilitates reconciliation and should therefore have a positive effect on fertility. However, part-time work also favors a gender specialization model in which women are secondary earners and main caregivers. On the other hand, Southern Europe (Spain, Portugal) have precarious working conditions. This is an especially acute problem in Southern Europe where it is not atypical for school graduates to wait two or three years before entering into a stable employment relationship.

The results for these four countries are consistent with Cazzola et al. (2016), Raymo and Shibata (2017) and others that also found a negative relationship between fertility and unemployment. However, the results for the other countries seem to indicate that there is no apparent relation between the two variables.

6. Conclusions

The relationship between having good employment and deciding to have more children is one which has been studied many times. In this paper we have examined the relationship between the unemployment rate and the fertility rate using fractional integration and cointegration methods for the case of ten countries of the EU along with

Japan and the US. We do not find any evidence of a long run equilibrium relationship between these two variables in any of the countries examined by using cointegration methods. However, assuming that the previous value of unemployment is weakly exogenous in the explanation of fertility, it produces a significant negative effect on current fertility rates at least in four of the countries examined in this work, namely, the Netherlands, Portugal, Spain and the US.

It should be noted that the empirical results in this study are subject to some limitations with regard to the data time span and omitted variable bias. Thus, further research should be conducted on this topic. In this context, the fractional cointegrated VAR (FCVAR) approach developed by Johansen and Nielsen (2010, 2012) can also be implemented with these two variables. Moreover, other variables that might affect fertility rates can be included in a regression model with long memory errors. Also, the possibility of non-linear structures within the fractional integration framework is another interesting issue that deserves to be investigated. Note that fractional integration and structural breaks are issues which are linked in many ways (Diebold and Inoue, 2001; Granger and Hyung, 2004; Ohanissian et al., 2008; Aue and Horváth, 2013; etc.) and though there exist procedures for testing this hypothesis (e.g., Gil-Alana, 2008) the limited number of observations invalidates its performance in the present context. Nevertheless, work in all these directions will be considered in future papers.

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Figure 1: Time series plots:Fertility rates

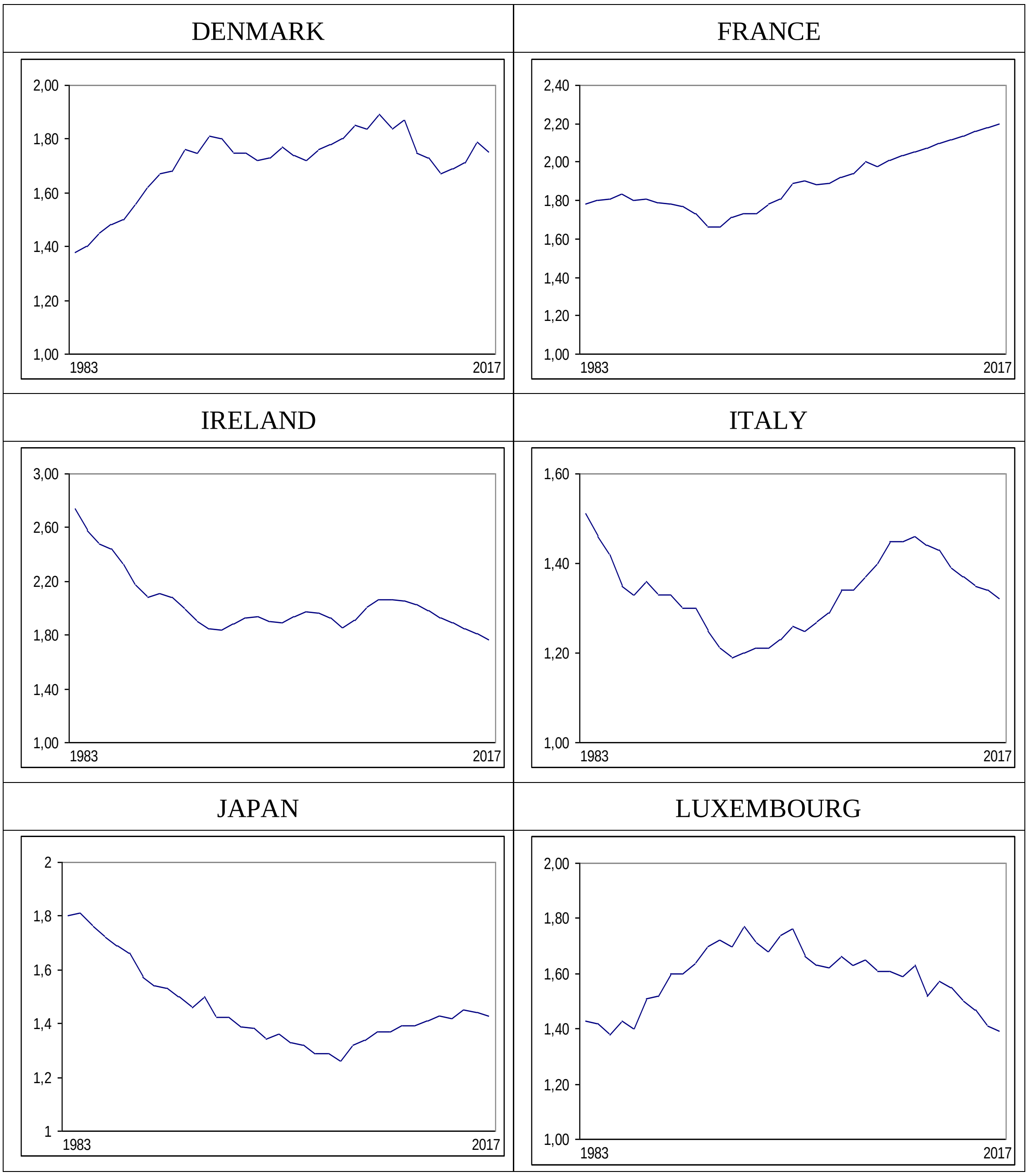


Figure 1: Time series plots: Fertility rates (cont.)

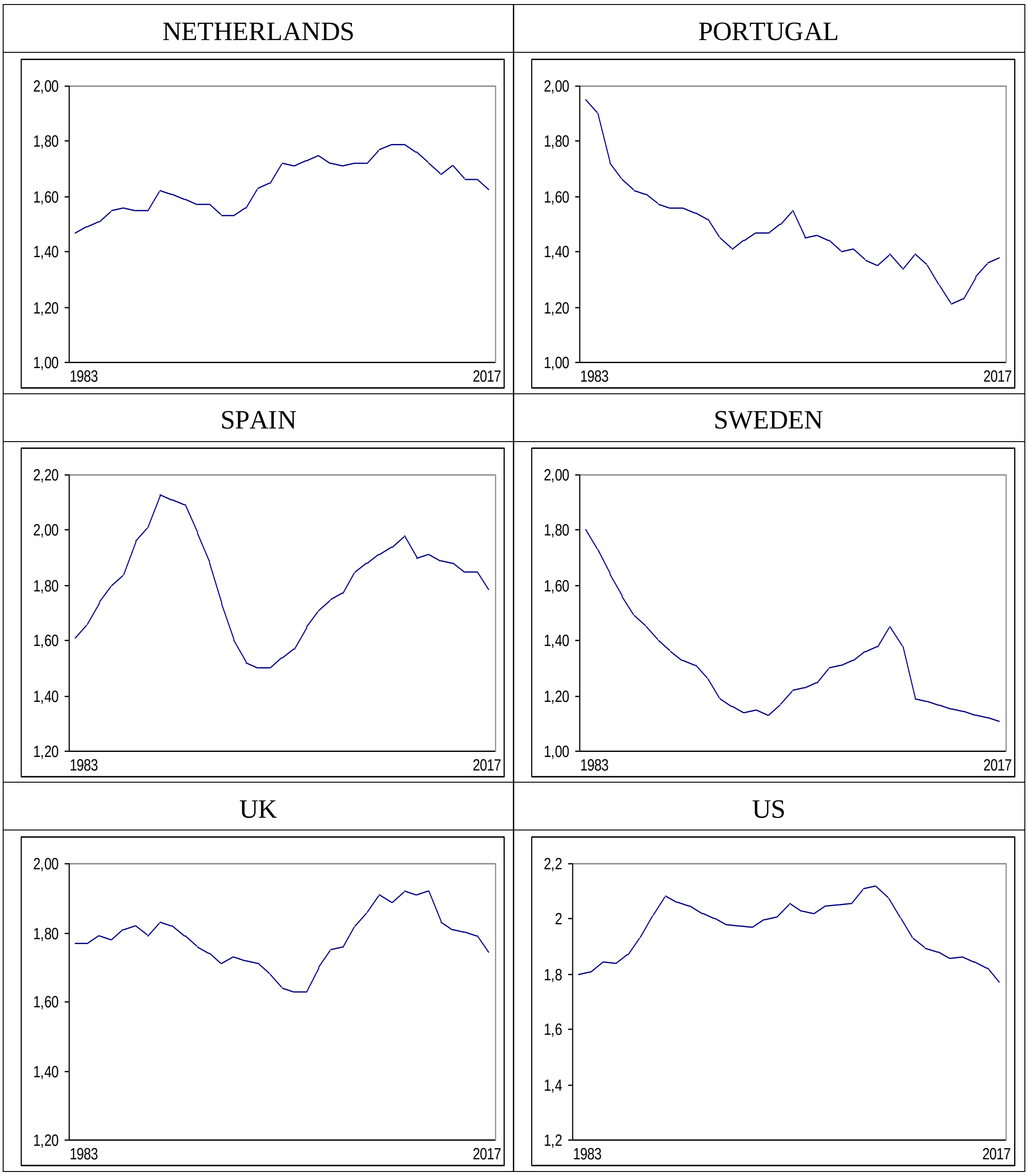
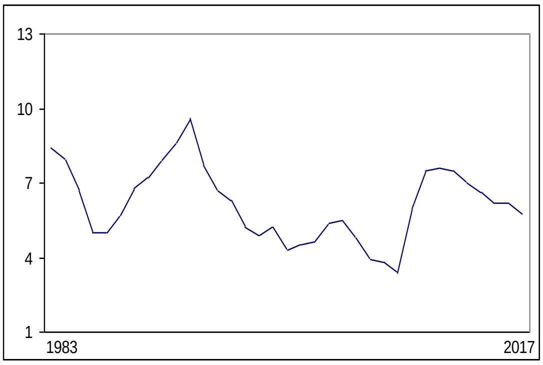
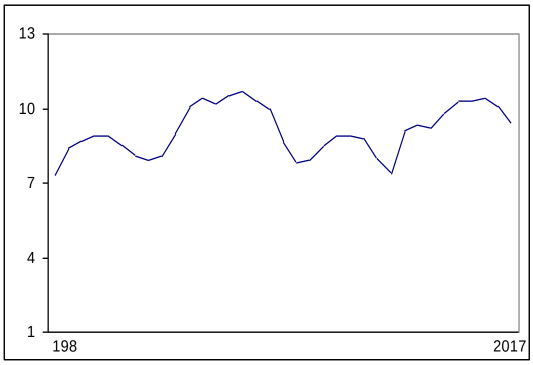


Figure 2: Time series plots: Unemployment rates

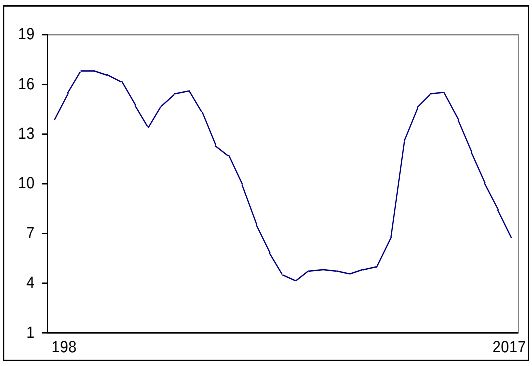
DENMARK



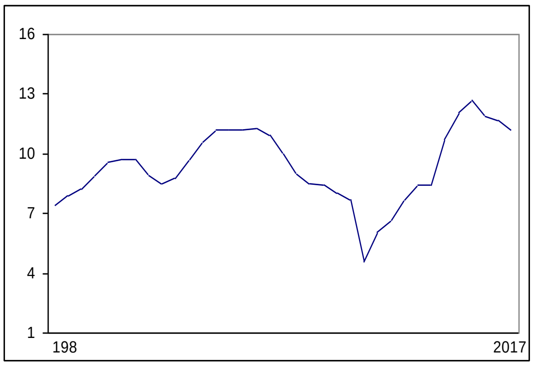
FRANCE



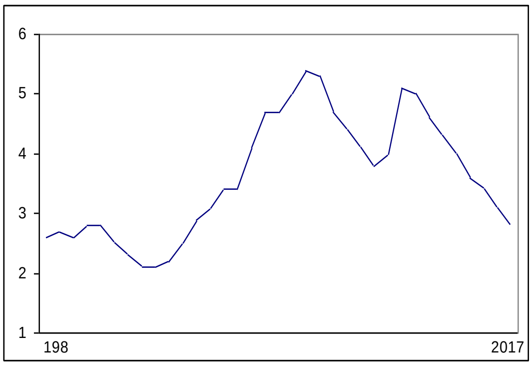
IRELAND



ITALY



JAPAN



LUXEMBOURG

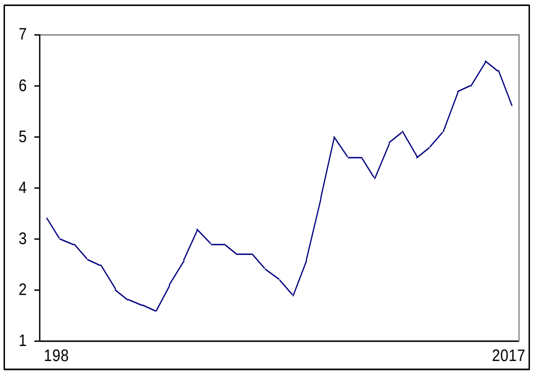


Figure 2: Time series plots: Unemployment rates (cont.)

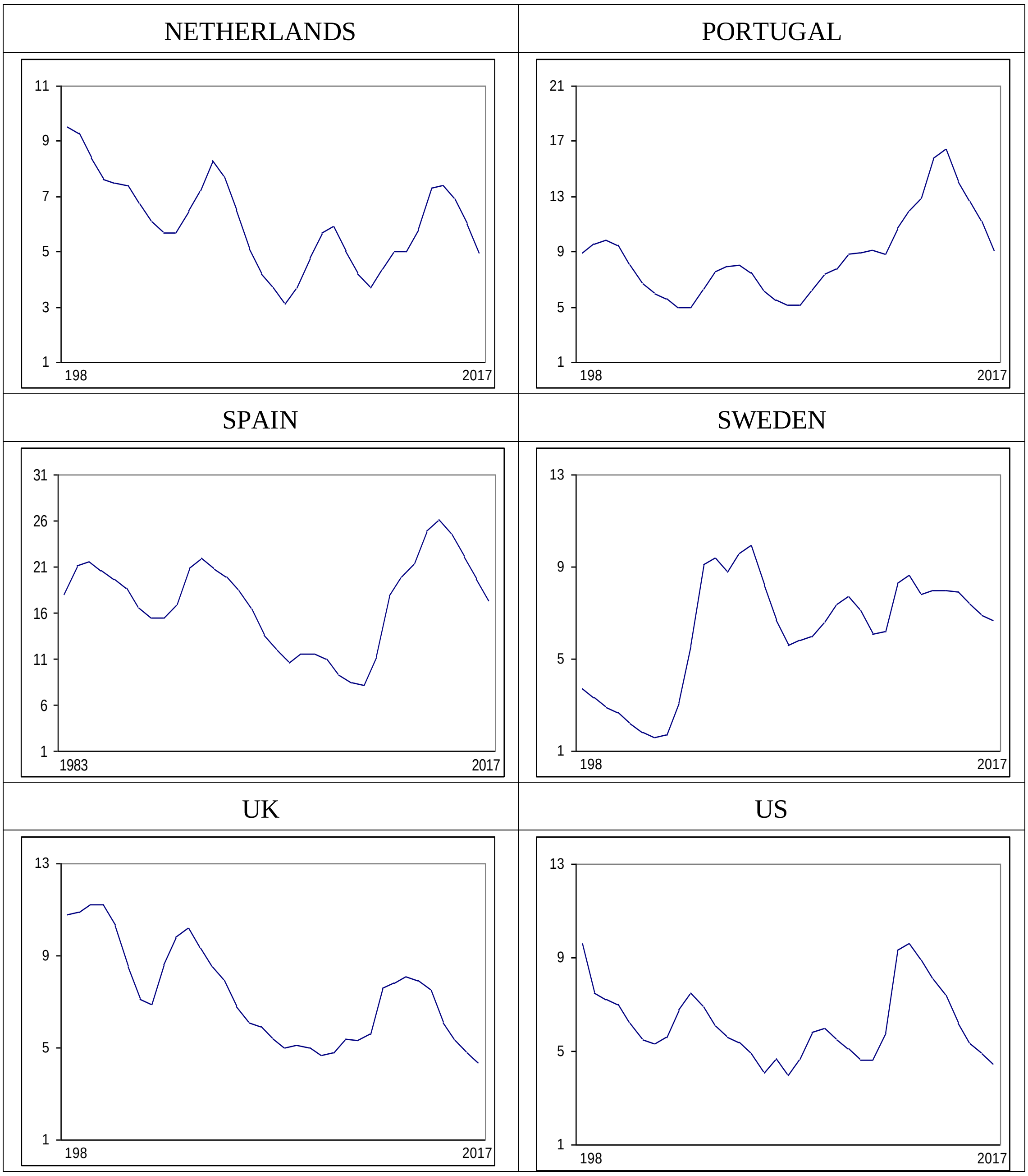


Table 1: Descriptive statistics

i) Fertility rate					
Country	Max.	Min.	Range	Mean	Std. Dev.
DENMARK	1,89	1,38	0,51	1,71	0,1296
FRANCE	2,02	1,66	0,36	1,86	0,10647
IRELAND	2,58	1,77	0,81	2,03	0,2242
ITALY	1,51	1,19	0,32	1,33	0,0862
JAPAN	1,81	1,26	0,55	1,46	0,1489
LUXEMBOURG	1,77	1,38	0,39	1,58	0,11398
NETHERLANDS	1,79	1,47	0,32	1,64	0,09223
PORTUGAL	1,95	1,21	0,74	1,47	0,1614
SPAIN	1,49	1,11	0,38	1,26	0,1127
SWEDEN	2,13	1,5	0,63	6,24	2,4787
U.K.	1,92	1,63	0,29	1,78	0,07781
U.S.A.	2,12	1,76	0,36	1,96	0,09983
ii) Unemployment rate					
Country	Max.	Min.	Range	Mean	Std. Dev.
DENMARK	9,6	3,4	6,2	6,15	1,4977
FRANCE	10,7	7,4	3,3	9,11	0,98
IRELAND	16,8	4,2	12,6	10,97	4,5727
ITALY	12,7	4,6	8,1	9,36	1,8356
JAPAN	5,4	2,1	3,3	3,6	1,0193
LUXEMBOURG	6,5	1,6	4,9	3,62	1,4706
NETHERLANDS	9,5	3,1	6,4	6,05	1,63371
PORTUGAL	16,4	5	11,4	8,71	2,9991
SPAIN	26,1	8,2	17,9	16,84	5,0295
SWEDEN	9,9	1,6	8,3	1,81	0,1742
U.K.	11,2	4,3	6,9	7,31	2,1467
U.S.A.	9,6	4	5,6	5,96	1,5524

Table 2: Estimated values of d on the first seasonal differences

i) Fertility			
Series	No terms	With intercept	With time trend
DENMARK	0.92 (0.70, 1.21)	**1.06 (0.85, 1.30)**	1.06 (0.89, 1.28)
FRANCE	0.86 (0.60, 1.20)	**1.23 (1.00, 1.53)**	1.23 (1.00, 1.54)
IRELAND	0.80 (0.55, 1.15)	1.65 (1.36, 2.42)	**1.55 (1.31, 2.49)**
ITALY	0.81 (0.56, 1.16)	**1.44 (1.26, 1.69)**	1.41 (1.25, 1.63)
JAPAN	0.90 (0.68, 1.23)	**1.13 (1.00, 1.31)**	1.11 (1.00, 1.27)
LUXEMBOURG	0.86 (0.63, 1.17)	**0.97 (0.84, 1.15)**	0.97 (0.84, 1.15)
NETHERLANDS	0.87 (0.62, 1.19)	**1.10 (0.86, 1.42)**	1.10 (0.86, 1.42)
PORTUGAL	0.87 (0.63, 1.22)	1.08 (0.51, 1.51)	**1.07 (0.78, 1.42)**
SPAIN	0.83 (0.58, 1.19)	**1.43 (1.19, 1.82)**	1.38 (1.17, 1.73)
SWEDEN	0.96 (0.76, 1.25)	**1.66 (1.43, 1.92)**	1.66 (1.43, 1.92)
U.K.	0.86 (0.63, 1.19)	**1.25 (1.03, 1.54)**	1.25 (1.03, 1.54)
U.S.A.	0.89 (0.67, 1.18)	**1.46 (1.19, 1.92)**	1.46 (1.19, 1.92)
ii) Unemployment			
Series	No terms	With intercept	With time trend
DENMARK	0.80 (0.56, 1.22)	**1.29 (0.82, 1.82)**	1.29 (0.85, 1.80)
FRANCE	0.97 (0.74, 1.31)	**1.13 (0.68, 1.77)**	1.13 (0.72, 1.89)
IRELAND	1.09 (0.84, 1.48)	**1.70 (1.35, 2.18)**	1.72 (1.35, 2.17)
ITALY	1.01 (0.82, 1.29)	**1.16 (0.91, 1.51)**	1.16 (0.92, 1.52)
JAPAN	0.94 (0.62, 1.37)	**1.36 (1.10, 1.82)**	1.36 (1.10, 1.82)
LUXEMBOURG	0.72 (0.55, 1.06)	**1.23 (0.89, 1.81)**	1.23 (0.84, 1.76)
NETHERLANDS	0.80 (0.57, 1.20)	**1.86 (1.22, 2.51)**	1.82 (1.19, 2.52)
PORTUGAL	0.84 (0.56, 1.34)	**1.64 (1.27, 2.10)**	1.64 (1.26, 2.13)
SPAIN	1.07 (0.79, 1.53)	**1.71 (1.27, 2.28)**	1.79 (1.28, 2.36)
SWEDEN	1.17 (0.77, 1.73)	**1.49 (1.08, 2.18)**	1.50 (1.07, 2.18)
U.K.	0.91 (0.64, 1.31)	**1.57 (1.08, 2.20)**	1.56 (1.08, 2.20)
U.S.A.	0.62 (0.28, 1.05)	**1.37 (0.93, 1.92)**	1.36 (0.95, 2.01)

The values in parenthesis indicate the 95% confidence band for the values of d; in bold the selected model for each series.

Table 3: Confidence intervals for the values of d

Country	Fertility rate	Unemployment rate
DENMARK	(0.85, 1.30)	(0.82, 1.82)
FRANCE	(1.00, 1.53)	(0.68, 1.77)
IRELAND	(1.31, 2.49)	(1.35, 2.18)
ITALY	(1.26, 1.69)	(0.91, 1.51)
JAPAN	(1.00, 1.31)	(1.10, 1.82)
LUXEMBOURG	(0.84, 1.15)	(0.89, 1.81)
NETHERLANDS	(0.86, 1.42)	(1.22, 2.51)
PORTUGAL	(0.78, 1.42)	(1.27, 2.10)
SPAIN	(1.19, 1.82)	(1.27, 2.28)
SWEDEN	(1.43, 1.92)	(1.08, 2.18)
U.K.	(1.03, 1.54)	(1.08, 2.20)
U.S.A.	(1.19, 1.92)	(0.93, 1.92)

Table 4: Estimated coefficients based on a fractional cointegrated approach

i) Fertility / Unemployment			
Country	d	Intercept (γ_0)	Slope coefficient (γ_1)
DENMARK	1.08 (0.86, 1.34)	**1.4261 (17.14)**	-0.0062 (-0.73)
FRANCE	1.15 (0.98, 1.42)	**1.8766 (30.84)**	**-0.0141 (-1.83)**
IRELAND	1.71 (1.37, 2.41)	**2.9013 (30.34)**	-0.0066 (-1.01)
ITALY	1.42 (1.22, 1.66)	**1.5715 (39.48)**	-0.0057 (-1.24)
JAPAN	1.13 (0.97, 1.30)	**1.8122 (34.40)**	-0.0025 (-0.15)
LUXEMBOURG	0.97 (0.81, 1.15)	**1.4252 (18.72)**	0.0015 (0.08)
NETHERLANDS	0.95 (0.70, 1.26)	**1.6516 (25.29)**	**-0.0189 (-3.03)**
PORTUGAL	1.11 (0.46, 1.42)	**2.1187 (26.42)**	**-0.0171 (-2.34)**
SPAIN	1.39 (1.15, 1.80)	**1.9025 (23.59)**	-0.0040 (-1.00)
SWEDEN	1.65 (1.42, 1.90)	**1.6049 (30.11)**	-0.0054 (-0.62)
U.K.	1.26 (0.98, 1.60)	**1.7804 (21.14)**	-0.0108 (-0.14)
U.S.A.	1.48 (1.24, 1.83)	**1.9255 (31.46)**	**-0.0138 (-2.58)**
i) Unemployment / Fertility			
Country	d	Intercept	Slope coefficient
DENMARK	1.35 (0.84, 1.95)	**12.9618 (2.91)**	-3.1516 (-0.98)
FRANCE	1.05 (0.63, 1.72)	**17.5900 (3.01)**	**-5.8208 (-1.78)**
IRELAND	1.75 (1.38, 2.21)	**25.6891 (2.16)**	-4.4469 (-1.05)
ITALY	1.04 (0.77, 1.45)	**22.8623 (2.92)**	**-10.2407 (-1-99)**
JAPAN	1.37 (1.09, 1.78)	2.2304 (0.73)	0.1913 (0.11)
LUXEMBOURG	1.24 (0.88, 1.79)	2.4694 (1.10)	0.7183 (0.46)
NETHERLANDS	1.76 (0.44, 2.47)	**16.5758 (3.78)**	-4.7549 (-1.59)
PORTUGAL	1.57 (1.25, 2.01)	**19.2469 (3.19)**	**-5.3350 (-1.77)**
SPAIN	1.68 (1.20, 2.29)	**22.5941 (1.87)**	-3.1778 (-0.48)
SWEDEN	1.47 (1.03, 2.16)	7.5955 (1.51)	-2.3613 (-0.75)
U.K.	1.67 (0.91, 2.34)	**17.3903 (2.79)**	-3.7447 (-1.07)
U.S.A.	1.41 (1.06, 1.84)	**30.4379 (3.83)**	**-11.3198 (-2.55)**

In bold, significant coefficients at the 5% level. In parenthesis in 3rd and 4rd columns, t-values.

Table 5: Estimated coefficients based on a fractional integrated approach

i) Fertility / Unemployment(-1) with white noise errors			
Country	d	Intercept (δ_0)	Slope coefficient (δ_1)
DENMARK	1.08 (0.82, 1.37)	**1.3425 (15.85)**	0.0058 (0.68)
FRANCE	1.18 (1.01, 1.48)	**1.8049 (27.98)**	-0.0009 (-0.12)
IRELAND	1.64 (1.33, 2.23)	**2.7134 (29.89)**	-0.0066 (-1.06)
ITALY	1.39 (1.14, 1.69)	**1.4907 (36.46)**	-0.0018 (-0.40)
JAPAN	1.13 (0.97, 1.33)	**1.8415 (34.62)**	-0.0076 (-0.46)
LUXEMBOURG	0.97 (0.77, 1.14)	**1.4251 (18.12)**	-0.0013 (-0.07)
NETHERLANDS	0.77 (0.55, 1.17)	**1.7336 (29.37)**	**-0.0246 (-4.20)**
PORTUGAL	1.03 (0.42, 1.40)	**2.0826 (25.20)**	**-0.0197 (-2.59)**
SPAIN	1.51 (1.27, 1.86)	**2.0469 (32.46)**	**-0.0163 (-5.02)**
SWEDEN	1.61 (1.32, 1.92)	**1.6425 (29.80)**	-0.0048 (-0.53)
U.K.	1.24 (1.00, 1.54)	**1.7268 (20.30)**	0.0036 (0.49)
U.S.A.	1.35 (1.11, 1.76)	**1.9069 (29.30)**	**-0.0118 (-2.06)**
i) Unemployment / Fertility			
Country	d	Intercept	Slope coefficient
DENMARK	1.33 (-0.57, 1.94)	**1.28859 (15.42)**	0.0107 (1.26)
FRANCE	1.03 (0.73, 1.46)	**1.8076 (28.53)**	-0.0011 (-0.14)
IRELAND	0.88 (0.36, 1.32)	**2.5569 (31.41)**	-0.0011 (-0.24)
ITALY	1.52 (1.04, 1.96)	**1.4815 (37.53)**	-0.0003 (-0.07)
JAPAN	1.90 (1.43, 2.51)	**1.8266 (42.90)**	0.0023 (0.16)
LUXEMBOURG	1.41 (0.94, 1.93)	**1.3939 (18.84)**	0.0090 (0.52)
NETHERLANDS	0.73 (0.25, 1.50)	**1.7462 (31.21)**	**-0.0257 (-4.56)**
PORTUGAL	0.97 (0.49, 1.52)	**2.0652 (26.24)**	**-0.0194 (-2.71)**
SPAIN	1.32 (-0.38, 1.91)	**2.0330 (32.78)**	**-0.0155 (-5.06)**
SWEDEN	2.21 (-0.39, 2.46)	**1.5964 (36.18)**	0.0057 (0.72)
U.K.	1.14 (0.35, 1.82)	**1.7248 (20.61)**	0.0040 (0.56)
U.S.A.	1.04 (0.44, 1.54)	**1,9394 (30.65)**	**-0.0146 (-2.58)**

In bold, significant coefficients at the 5% level. In parenthesis in 3rd and 4rd columns, t-values.