On the Persistence of Labor Force Participation Rates by Gender: Evidence from OECD Countries

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Abstract

We present empirical evidence regarding differences in the time series properties of labour force participation rates across gender for a group of twelve OECD countries. Our results indicate that there are gender differences in the dynamics of labor force participation rates across countries. Specifically, the female labor force participation rates are relatively more persistent in seven countries (Australia, Canada, Germany, Japan, The Netherlands, Portugal, and Spain), and the male labour force participation rate is more persistent in four countries (Finland, Norway, Sweden, and the U.S.).

Keywords: Labour Force Participation Rates, Gender Gap, Persistence, Unit Root

JEL Classification: E24

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1. Introduction

The use of the unemployment rate to ascertain the strength or state of the labour market has recently been brought into question. Murphy and Topel (1997) and Krugman (2004), for instance, have argued that there has been a decrease in the labour force participation rates in the United States that point to a weakening of its labour market. There are two main arguments for why the unemployment rate may not, by itself, be a reliable indicator of the state of the labour market. First, the possible presence of a 'discouraged worker' effect implies that during recessionary periods, individuals may drop out of the labour market and are, therefore, not classified as either employed or unemployed. In the presence of a discouraged worker effect, then, the unemployment rate under-estimates the extent of joblessness in the economy, but the decrease in the labour force participation rate will provide insight into the weakening of the labour market. On the other hand, during expansionary periods, the 'added worker' effect that reflects an increase in the labour supply of married women owing to a loss of their husbands jobs, may result in a decrease of the unemployment rate but an increase in the labour force participation rate. Therefore, use of the labour force participation rate along with the unemployment rate can provide more insight regarding the relative strength of the labour market.

Second, as argued by Gustavsson and Österholm (2006, 2010), Madsen, Mishra, and Smyth (2008), and Ozdemir, Balcilar, and Tansel (2013), the informational value of the unemployment rate depends on the time series properties of the labour force participation rate. The unemployment rate (u) and the labour force participation rate (LFPR) are defined as:

$$u = \frac{U}{(U+E)}$$

$$LFPR = \frac{(U+E)}{POP}$$

where U is the number of unemployed, E is the number of employed, and POP is the total population. An increase in the unemployment rate may be the result of an increase in U or a decrease in E. If a change in E is matched one-for-one by a corresponding change in U, then changes in the unemployment rate will reflect the state of the labour market since the labour force participation rate remains constant or mean reverting.

Consider the implication of a non-stationary labour force participation rate. In this case, any shock to the labour force participation rate will result in a permanent or persistent shift in the participation rate. And, a change in E will not correspond to a one-for-one change in U. If an observed increase in the unemployment rate is due to an increase in the unemployed (U), then the labour force participation rate will increase. But, if the observed increase in the unemployment rate is due to a decrease in the employed (E), then the labour force participation rate will decrease. Therefore, we cannot ascertain the implications regarding the state of the labour market but just looking at the unemployment rate.

In recent years, there has been a growing literature regarding the persistence of unemployment rates by gender, see Ewing, Levernier, and Malik (2005), Queneau and Sen (2008), and Queneau and Sen (2012). Although many studies have looked at differences in labor supply between men and women, very few have examined the degree of persistence of labor force participation rates by gender across countries. Two exceptions to this are Gustavsson and Österholm (2010) who examined the labour force participation rates by gender for the U.S., and Ozdemir, Balcilar, and Tansel (2013) who examined the labour force participation rates by gender for Australia, Canada, and the U.S. The evidence presented in these two studies is mixed. While Gustavsson and Österholm (2010) do not find evidence of stationarity, Ozdemir, Balcilar, and Tansel (2013) do find evidence of stationarity.[1]

The objective of our paper is to examine any differences in the time series properties of the labour force participation rates between men and women as well as across countries based on a sample of twelve OECD countries. Our evidence indicates that the labour force participation rates for men are stationary in all countries except Canada, and that the labour force participation rates for women are stationary in Australia, Finland, Germany, Japan, The Netherlands, Portugal, and Spain. In Canada, Italy, Norway, Sweden, the U.S., the female labour force participation rates are non-stationary.

The rest of our paper is structured as follows. In Section 2, we discuss the pattern of labour force participation rates across countries and across gender. We also provide

[1] To the best of our knowledge, there are also very few studies that examine the time series properties of the aggregate labor force participation rates. Gustavsson & Österholm (2006) examined the labor force participation rates in Australia, Canada, and the U.S. and found them to be non-stationary. Madsen, Mishra, and Smyth (2008) examined the labour force participation rates of G7 countries, and found evidence of non-stationarity in France, Italy, and the U.S. However, using fractionally integrated models with endogenous structural breaks, Ozdemir, Balcilar, and Tansel (2013) examined the labor force participation rates in Australia, Canada, and the U.S., and found these to be stationary.

some insight regarding some factors that may contribute to gender differences in the dynamics of labor force participation rates across countries. We present empirical evidence regarding the time series properties of the labour force participation rates in Section 3. In Section 4, we provide some comments regarding the policy implications of our findings.

2. Dynamics of Labour Force Participation Rates

In this section, we discuss the differences in the labour force participation rates by gender across twelve OECD countries: Australia, Canada, Finland, Germany, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, and the United States using annual data obtained from the OECD. The group of countries in our sample form interesting sub-groups, namely, South-Europe (Italy, Spain, and Portugal), North-Europe (Germany and the Netherlands), Scandinavian (Finland, Norway, Sweden), and Anglo (Australia, Canda, and the United States). Table 1 presents the sample period for which data were available for each country, and Figures 1-12 show the evolution of the male and female labour force participation rates for each country in our sample.

There are some interesting patterns regarding the evolution of the labour force participation rates across countries and across gender. The female labour force participation rates ($LFPR^F$) are lower compared to the respective male labour force participation rates ($LFPR^M$) across all countries. The $LFPR^M$ series tend to fall over the respective sample period except in Japan and the Netherlands, but the $LFPR^F$ series increase over the respective sample period across all countries. In Finland, Norway, Sweden, and the U.S., the female labour force participation rates seem to level off around the mid to late 1980's. In all other countries, there appears to be a gradual increase in the female labour force participation rates over the entire sample period. In Italy, Spain, and the Netherlands, the female labour force participation rates have the most dramatic increase, while in Finland, the increase is the least.

Over the last thirty years, male and female labor force participation rates have become closer in the OECD countries. However, gender is still an important source of worker heterogeneity in these countries. There are traditional supply-side factors such as gender differences in human capital accumulation and job search behaviors that may explain differences the level and dynamics of male and female labor force participation rates. Institutional factors such as unemployment benefits, mandatory family benefits, and the extent of employment discrimination against women may also explain gender differences in the persistence of labor force participation rates. Family leave policies may interact with social norms to change the labor participation of women. For example, women may take a longer pregnancy leave and, therefore, delay their re-entry into the labor market when labor market conditions are difficult because they know that it will be difficult to find employment. If there is employment discrimination against women there may feedback effects, see Arrow (1973), Blau, Ferber, and Winkler (2002). Women may postpone their decision to supply labor or even drop out of the labor force if they believe that the extent of employment discrimination against women makes the chance of getting a job is very small. Put differently, there is discouraged worker effect if women are discriminated against in the labor market. As pointed out by Blau, Ferber, and Winkler (2002):

> "*Even a relatively small amount of initial labor market discrimination can have greatly magnified effects if it discourages women from making human capital investments, weakens their attachment to the labor force, and provides economic incentives for the family to place priority on the husband's career*".

We focus on differences in the dynamics of labour force participation rates across gender and across countries. Understanding the time series properties of labour force participation rates by gender are important for the several reasons. First, the labour force participation rate is in itself an important indicator of labor market performance and gender equality in the labor market. Second, differences in the level of persistence of labor force participation rates across gender imply that shocks impact differently and female labor force participation rate compared to the male labour force participation rate. Therefore, policy makers should take into account the extent of differences in the level of persistence in the labour force participation rate across men and women when devising policies, for instance, during recessionary periods. Finally, as previously discussed, the dynamics of labor participation rates impacts the informational value of the unemployment rate. If for example female labor participation rates tend to be more persistent than male labor force participation rates then the unemployment rate is a less accurate measure of the strength of the labour market for women than for men.

In order to analyze the time series properties of labour force participation rates, we propose using the three established theories of unemployment, namely, the natural rate theory, the structural approach, and the hysteresis hypothesis. If the labour force participation rate evolves according to the natural rate theory, we would view it as mean reverting, that is, the labour force participation rate would be characterized as transitory shocks around a constant mean. The structural approach implies that the labour force participation rate is stable around a trend or that it has a structural break in the mean and/or trend. In this case, shocks to the labour force participation rate have a transitory effect albeit around a trend or changing mean/trend. Within the natural rate of structural paradigms, the level of persistence in the labour force participation rates is relatively low. The hysteresis hypothesis, however, implies that the labour force participation rate has a unit root, and so any shock to it will have a persistent effect. Knowledge of the level of persistence in the labour force participation rate can be useful in gauging the extent and duration of the effect of specific policies that impact the labour force participation rate.

In order to assess the dynamics of the labour force participation rate series, we test for the presence of a unit root in these series. The appropriate characterization that describes the dynamics of the labour force participation rate depends on whether the series has a unit root or not. If the unit root hypothesis is not rejected, the appropriate characterization of the labour force participation rate is the hysteresis hypothesis, and so the level of persistence in the series is relatively high. As well, in this case, the unemployment rate is not very informative regarding the state of the labour market. If, however, the unit root hypothesis is rejected, then we would infer that the labour force participation rate evolves according to the natural rate theory or the structural approach, and that the level of persistence in the series is relative low. In this case, the particular characterization of the dynamics of the labor force participation rates depends on whether there is a trend and/or a structural break. If there is no trend then the appropriate characterization of the labour force participation rate is the natural rate theory. If, on the other hand, there is a trend or a structural break in the mean and/or trend, we would use the structural approach to characterize the labour force participation rate.

3. Empirical Results

In this section, we test for the presence of a unit root in the $LFPR^F$ and $LFPR^M$ series of all countries in our sample. We use different versions of unit root tests to determine the appropriate characterization of labour force participation rate dynamics across gender and across countries. In addition, we compare the level of persistence between the male and female labour force participation rates using the measure of half-life, see Andrews (1993) for details.

First, calculate the Augmented Dickey-Fuller (ADF) unit root tests using the following regressions:

$$y_t = \hat{\mu} + \hat{\rho}\, y_{t-1} + \sum_{j=1}^{k^*} \hat{c}_j\, \Delta y_{t-j} + \hat{e}_t \tag{1}$$

$$y_t = \hat{\mu} + \hat{\beta}\, t + \hat{\rho}\, y_{t-1} + \sum_{j=1}^{k^*} \hat{c}_j\, \Delta y_{t-j} + \hat{e}_t \tag{2}$$

The ADF test from regression (1) without a time trend is denoted by τ_μ, and the ADF test from regression (2) with a time trend is denoted by τ_τ. Given that the $LFPR^F$ series exhibit a trend in all countries, we we did not calculate τ_μ statistic for these series. The calculated τ_μ, and τ_τ statistics for the $LFPR^M$ series, and the τ_τ statistics for the $LFPR^F$ series are summarized in Table 2. The τ_μ tests show that the $LFPR^M$ series for Australia, Germany, and Portugal are mean stationary, and the τ_τ show that the $LFPR^F$ series for Japan and Portugal are trend stationary.

We also use the unit root tests proposed by Popp (2008) and Costantini and Sen (2013) to allow for a one time break in the trend function at an unknown break-date.[2] Popp (2008) and Costantini and Sen (2013) use the conventional components representation of the underlying data generating process to test for the presence of unit root. We use their Model M_2, the Mixed Model, that allows for a simultaneous break in the intercept and slope of the underlying trend function. Their tests allow for the presence of a one-time break under both the unit root null and trend-break stationary alternative hypotheses. This unit root test is based on the following reduced form regression:

$$y_t = \alpha_2^* + \beta_2^* t + \kappa\, DU_{t-1}(T_B) + \xi\, D_t(T_B) + \zeta\, DT_{t-1}(T_B) + \rho\, y_{t-1} + \sum_{j=1}^{k^*} c_j \Delta y_{t-j} + e_t \tag{3}$$

where $DU_{t-1}(T_B) = 1_{(t > T_B + 1)}$, $DT_{t-1}(T_B) = (t - 1 - T_B)\, 1_{(t > T_B + 1)}$ and $D_t(T_B) = 1_{(t = T_B + 1)}$, 1(.) is the indicator function, and T_B is the break-date. The test regression (3) differs from the corresponding regression specifications of Perron (1997) since the intercept dummy $DU_{t-1}(T_B)$ and the trend dummy $DT_{t-1}(T_B)$ appear in their lagged forms, and the break parameter is no longer the coefficient of the dummy variable $DU_t(T_B)$ but the coefficient of the impulse dummy variable $D_t(T_B)$.

[2] Our data spans, at best, the period 1952-2011, and for most countries, data is available for an even shorter time period. In addition, the testing procedure of Popp (2008) and Costantini and Sen (2013) requires specification of the trimming parameter λ_0 (= 0.1) that reduces further the sample over which we search for a break in the trend function. Given that we view structural breaks as fundamental shifts in the economy, we decided to use the one-break unit root tests, and not multiple breaks.

When the true location of the break-date is unknown, regression (3) is estimated for all possible break-dates $T_B = [\lambda T]$ corresponding to λ in $[\lambda_0, 1 - \lambda_0]$. Popp (2008) suggests using the following estimated break-date:

$$\hat{T}_B = \arg\max_{T_B} \left| t_{\hat{\xi}}(T_B) \right| \tag{4}$$

Popp's (2008) unit root statistic, denoted by $t_{\hat{\rho}}^2$, is the t-statistic for $H_O: \rho = 1$ in regression (3) evaluated at the corresponding estimated break-date as defined in (4). In practice, the value of the lag-truncation parameter (k^*) is unknown, and so we use Perron and Voglesang's (1992) k(t-sig) method for selecting the lag-truncation parameter k^*.[3]

Costantini and Sen (2013) develop unit root tests for the joint null hypothesis of a unit root. Their tests are based on the fact that the reduced form estimation equation (3) imposed additional restriction under the unit root null hypothesis, that is, $\xi = 0$ in regression (3). Specifically, the joint null unit hypotheses for model M_2 is:

$$H_0^{M_2}: \beta_2^* = 0, \zeta = 0, \rho = 1 \tag{5}$$

For a given break-date T_B, the F-statistic $F_T^2(T_B)$ for the null hypothesis $H_0^{M_2}$ is defined as follows:

$$F_T^2(T_B) = \frac{\left(R_2\, \hat{\mu}_2(T_B) - r_2\right)' \left[R_2 \left\{ \sum_{t=1}^{T} x_T^2(T_B)\, x_T^2(T_B)' \right\}^{-1} R_2' \right]^{-1} \left(R_2\, \hat{\mu}_2(T_B) - r_2\right)}{3\hat{\sigma}_2^2(T_B)} \tag{6}$$

where $x_T^2(T_B)$ is the explanatory variables matrix, $\hat{\mu}_2(T_B)$ is the estimated parameter vector, and $\hat{\sigma}_2^2(T_B)$ is the estimated mean squared error from regression (3), $r_2 = (0, 0, 1)'$, and R_2 is the matrix corresponding to the null hypotheses $H_0^{M_2}$, that is:

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[3] First, we specify an upper bound `kmax' for the lag-truncation parameter. The chosen value of the lag-truncation parameter (k^*) is determined according to the following 'general-to-specific' procedure: the last lag in an autoregression of order k^* is significant, but the last lag in an autoregression of order greater than k^* is insignificant. The significance of the coefficient is assessed using the 10% critical values based on a standard Normal distribution.

$$R_2 = \begin{bmatrix} 0\,1\,0\,0\,0\,0\,0_k' \\ 0\,0\,0\,0\,1\,0\,0_k' \\ 0\,0\,0\,0\,0\,1\,0_k' \end{bmatrix}$$

Costantini and Sen's (2013) unit root test is the F-test for the joint null hypotheses $H_0^{M_2}$ evaluated at the estimated break-date defined in (4), that is, $F_T^2\left(\hat{T}_B\right)$.

The calculated Popp (2008) and Costantini and Sen (2013) statistics for all LFPR^{M} and LFPR^{F} series is given in Table 3. For each series, we report the test statistics, the estimated break-date, the estimated break-fraction, the estimated coefficient on the first lag of the dependent variable, and the estimated standard error. Based on the $t_{\hat{\rho}}^2$ statistic, we reject the unit root null hypothesis for the LFPR^{M} series in Finland, and for the LFPR^{F} series in Finland and Portugal. Based on the F_T^2 statistic, we reject the joint unit root null hypothesis for all LFPR^{M} series except Australia, Canada, Portugal, and the U.S., as well as, for the LFPR^{F} series in Australia, Finland, Germany, The Netherlands, Portugal, and Spain. For completeness, we also calculated Popp's (2008) and Costantini and Sen's (2013) unit root statistics corresponding to the Changing Mean Model (Model M_0) that allows for a break in the mean at an unknown date. We only calculated the Model M_0 statistics for the LFPR^{M} series as shown in Table 4. We did not test for the presence of a unit root using the Model M_0 unit root statistics in the LFPR^{F} series given that these series exhibit a clear trend for at least part of the sample. These results show that the LFPR^{M} series for Australia and the U.S. are stationary.[4]

Taken together, we find evidence of stationarity in all LFPR^{M} series except for Canada, and for the LFPR^{F} series in seven countries: Australia, Finland, Germany, Japan, The Netherlands, Portugal, and Spain. So, the LFPR^{M} series in Canada and the LFPR^{F} series in Canada, Italy, Norway, Sweden, the U.S. follow the hysteresis hypothesis.

We measure the degree of persistence in the labour force participation rate series using the half-life of a unit shock (HL_ρ). The half-lives are reported in Tables 2-4. We use the half-lives from Table 3 for our analysis given that these correspond to the most general model fitted to our series. The half-life, calculated as $|\log(1/2)/\log(\rho)|$, measures the time required for a shock to decay to half its initial value.[5] The half-lives for the male

[4] For a detailed discussion of Model M_0 and the corresponding unit root statistics, the reader is referred to Popp (2008) and Costantini and Sen (2013).

[5] See Andrews (1993) for a discussion of the half-life measure for persistence.

labour force participation rates range from 0.57 years for Norway to 6.60 years for Australia, and the female labour force participation rates range from 0.70 years for Portugal to 5.94 years for Sweden. In countries where both $LFPR^F$ and $LFPR^M$ are both stationary, the half-life indicates a lower level of persistence in the labour force participation rate for women compared to men except in Finland and the Netherlands. In countries where the $LFPR^M$ is stationary but the $LFPR^F$ series is not, the level of persistence among women is higher compared to men except in Italy where the half-life is approximately the same. In Canada where both the $LFPR^M$ and the $LFPR^F$ series is non-stationary, the level of persistence among men is higher than that of the women.

4. Policy Implications and Concluding Remarks

Based on our empirical evidence, there is more evidence of hysteresis in the labour force participation rates for women compared to men. More specifically, we find evidence of hysteresis in the $LFPR^F$ series in Canada, Italy, Norway, Sweden, and the U.S., and in the $LFPR^M$ series in Canda. Therefore, in these cases, policy makers should be cautious in using the unemployment rate as a measure of the strength of the labour market.

Further, our results indicate that the level of persistence in the labour force participation rate can be different for men and women across countries. In some instances, the level of persistence in the labour force participation rate is higher for women such as in Finland, Norway, Sweden, and the U.S. In Italy, the level of persistence in the labour force participation rates is almost the same, and for all other countries (Australia, Canada, Germany, Japan, The Netherlands, Portugal, and Spain), the level of persistence in the labour force participations is higher for men relative to women. From a policy perspective, this implies that care must be taken to ascertain the effects on the labour market of specific initiatives undertaken to stimulate the economy during recessions, or to determine the effects of shocks (positive or negative) to the economy that may impact the labour market. Consider, for instance, a negative shock to the economy. In countries like Finland, Norway, Sweden of the U.S., if the a shock to the economy causes a negative impact on the labour force participation rates, then we would expect the female labour force participation rate to take longer to return to its trend compared to the male labour force participation rate.

On the other hand, in countries where the level of persistence in the labour force participation rate for women is lower compared to men such as in Australia, Canada, Germany, Japan, The Netherlands, Portugal, and Spain, policy initiatives to stimulate the labour force participation rates will be more effective for men compared to women. Therefore, policy makers may want to consider tackling some of the more fundamental factors that impact the women's participation in the labour force such as the level of human capital accumulation, the job search behavior, institutional factors including unemployment benefits, mandatory family benefits, and the extent of employment discrimination against women.

References

Arrow, K. J., 1973, "The Theory of Discrimination", in O. Ashenfelter and A. Rees (eds.), *Discrimination in Labor Markets*, Princeton, NJ: Princeton University Press.

Andrews, D. W. K., 1993, "Exactly Median-Unbiased Estimation of First Order Autoregressive/ Unit Root Models," *Econometrica*, 61, 139-165.

Blau, F., Ferber, M., and Winkler, A., 2002, "Models of discrimination," from *The Economics of Women, Men and Work*, Fourth Edition, Prentice Hall: New Jersey, 219-235.

Costantini, M., and Sen, A., 2013a, "New Tests for the Joint Unit Root Null Hypothesis in Time Series with a Structural Break," Manuscript, Xavier University, Cincinnati, OH, U.S.A.

Costantini, M., and Sen, A., 2013b, "Critical Values for New Perron-Type Unit Root Tests," Manuscript, Xavier University, Cincinnati, OH, U.S.A.

Ewing, Levernier, and Malik, 2005, "Modeling unemployment rates by race and gender: A nonlinear time series approach," *Eastern Economic Journal,* 31, 333-47.

Gustavsson, M, and Osterhom, P., 2006, "The Informational Value of Unemployment Statistics: A Note on the Time Series Properties of Participation Rates," *Economics Letters*, 92 (3), 428-433.

Gustavsson, M., and Österhom, P., 2007, Does Unemployment Hysteresis Equal Employment Hysteresis?" *Economic Record*, 83, 159–173.

Krugman, P. R., 2004, "Checking the Facts, in Advance," *New York Times*, October 12.

Madsen, J., Mishra, V., and Symth, R., 2008, “Are Labour Force Participations Rates Non-Stationary? Evidence from 130 Years for G7 Countries,” *Australian Economic Papers*, 47 (2), 166-189.

Murphy, K. M., and Topel, R., 1997, “Unemployment and Nonemployment,” *American Economic Review*, 87, 295-300.

Osterholm, P., 2010, “Unemployment and Labour-Force Participation in Sweden,” *Economics Letters*, 106 (3), 205-208.

Ozdemir, Z. A., Balcilar, M., and Tansel, A., 2013, “International Labor Force Participation Rates by Gender: Unit Root or Structural Break?” *Bulletin of Economic Research*, 65 (1), 142-164.

Popp, S., 2008, “New Innovational Outlier Unit Root Test with a Break at an Unknown Time,” *Journal of Statistical Computation and Simulations*, 78, 1145-1161.

Queneau, H., and Sen, A., 2008, “Evidence on the Dynamics of Unemployment by Gender,” *Applied Economics*, 40 (16), 2099-2108.

Queneau, H., and Sen, A., 2012, “On the Structure of U.S. Unemployment Disaggregated by Race, Ethnicity, and Gender,” *Economics Letters*, 117 (1), 91-95.

Table 1: Countries and the Corresponding Time Periods

Country	Country Code	Period	Observations
Australia	AUS	1966 – 2011	46
Canada	CAN	1976 – 2011	36
Finland	FIN	1963 – 2011	49
Germany	DEU	1970 – 2011	42
Italy	ITA	1970 – 2011	42
Japan	JPN	1968 – 2011	44
Netherlands	NLD	1971 – 2011	41
Norway	NOR	1972 – 2011	40
Portugal	PRT	1974 – 2011	38
Spain	ESP	1972 – 2011	40
Sweden	SWE	1963 – 2011	49
United States	USA	1960 – 2011	52

Figure 1: $LFPR^M$ & $LFPR^F$ for Australia, 1966−2011

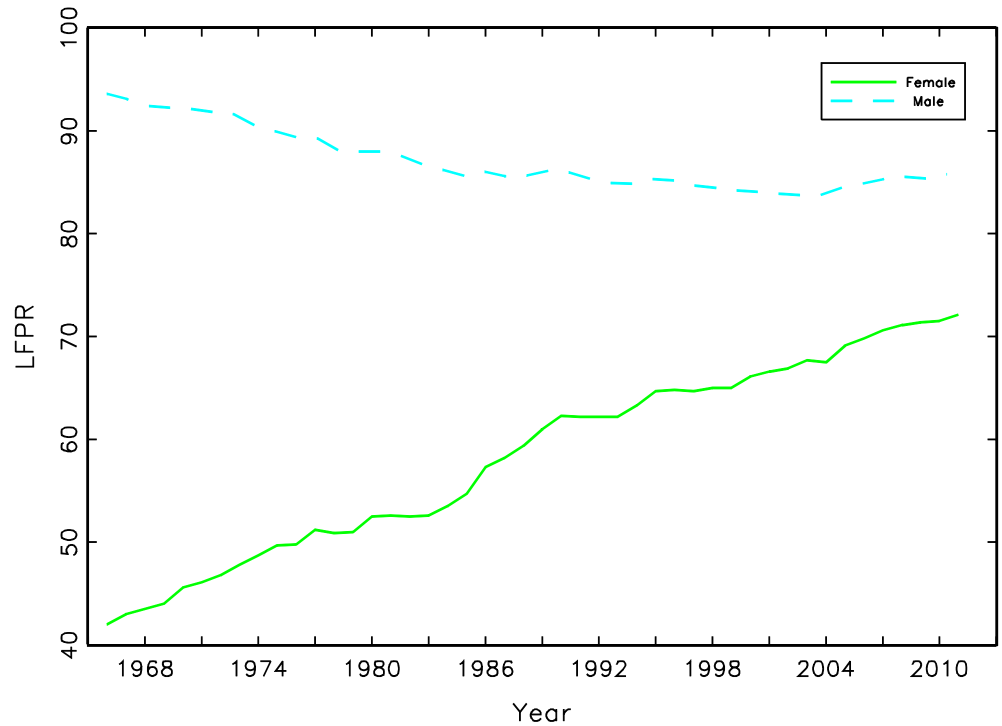

Figure 2: $LFPR^M$ & $LFPR^F$ for Canada, 1976−2011

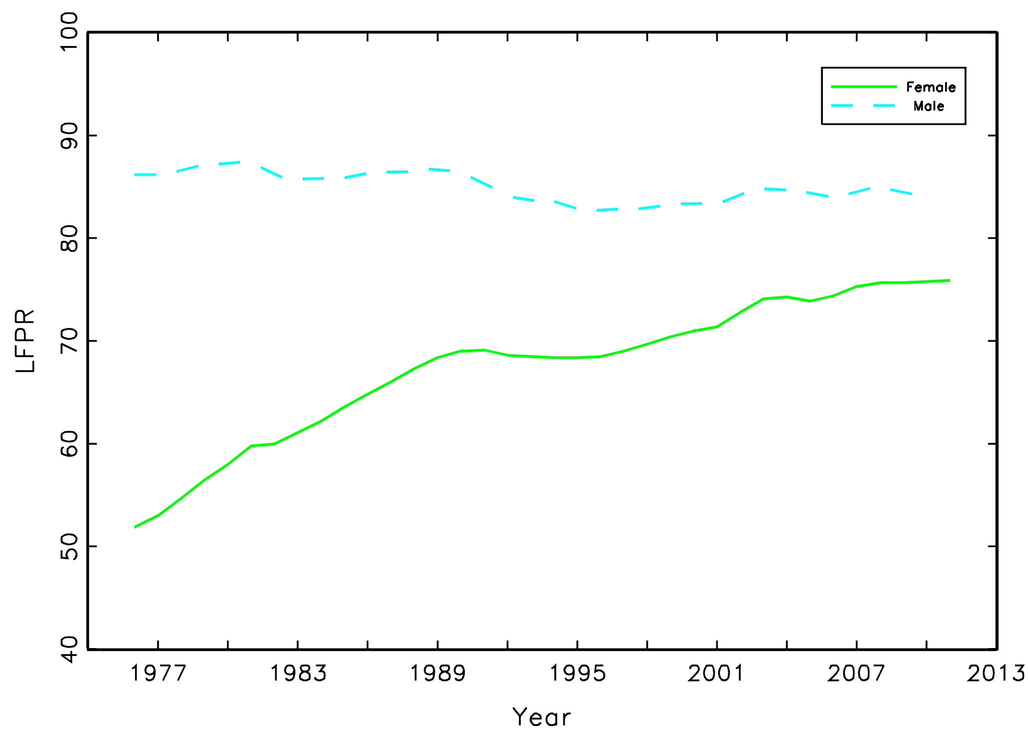

Figure 3: $LFPR^M$ & $LFPR^F$ for Finland, 1963–2002

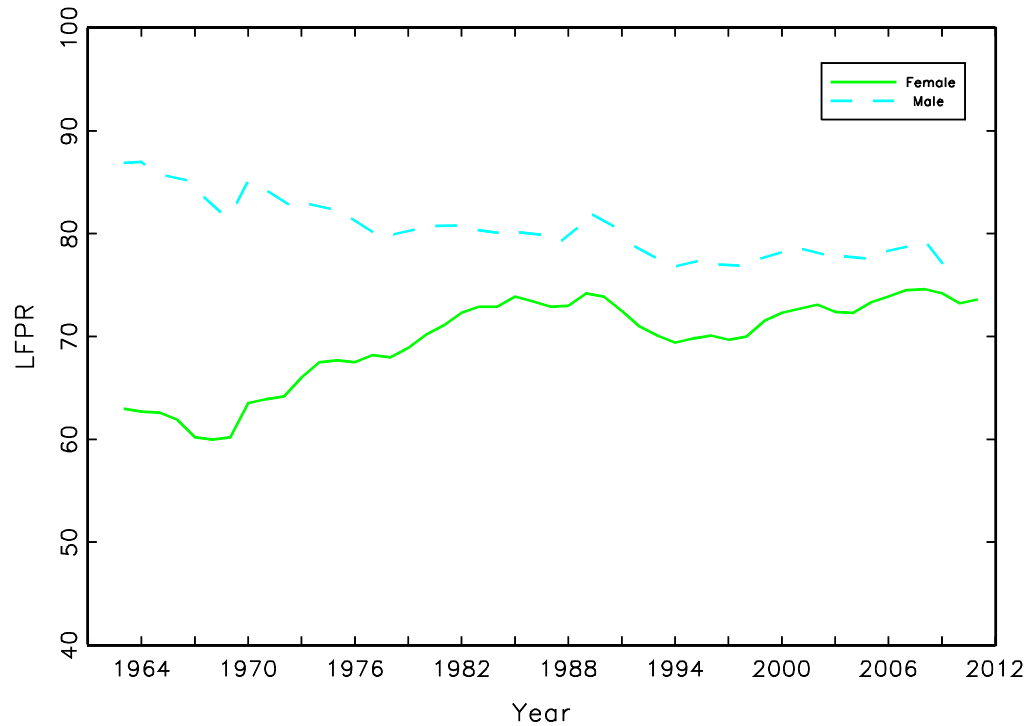

Figure 4: $LFPR^M$ & $LFPR^F$ for Germany, 1965–2002

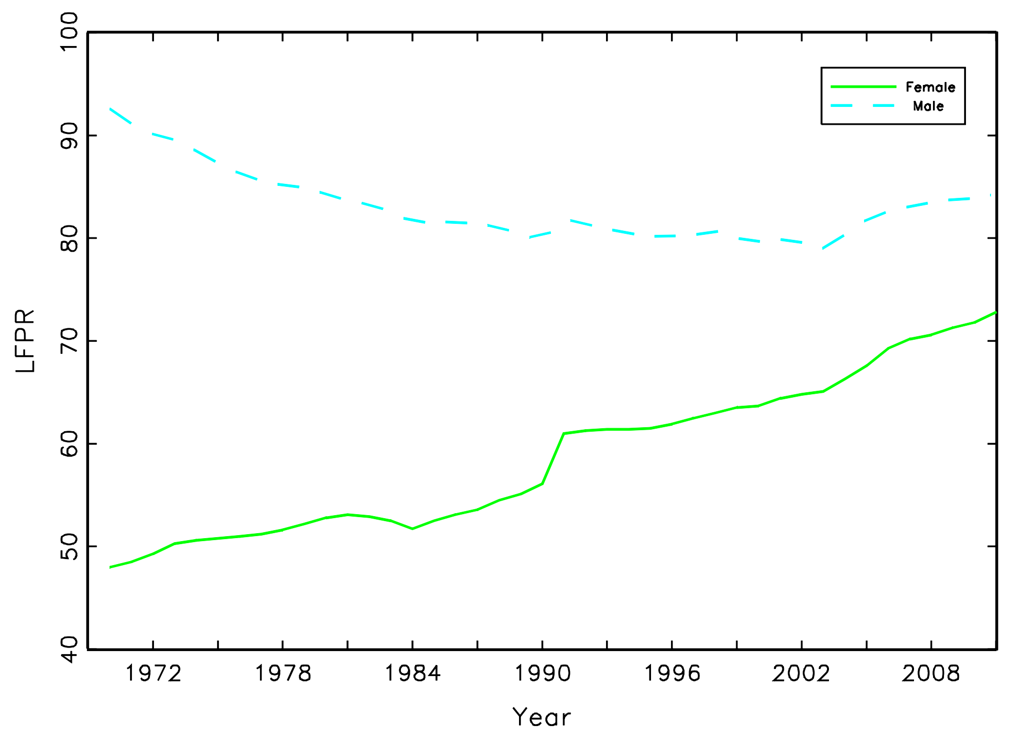

Figure 5: LFPR^M & LFPR^F for Italy, 1970–2002

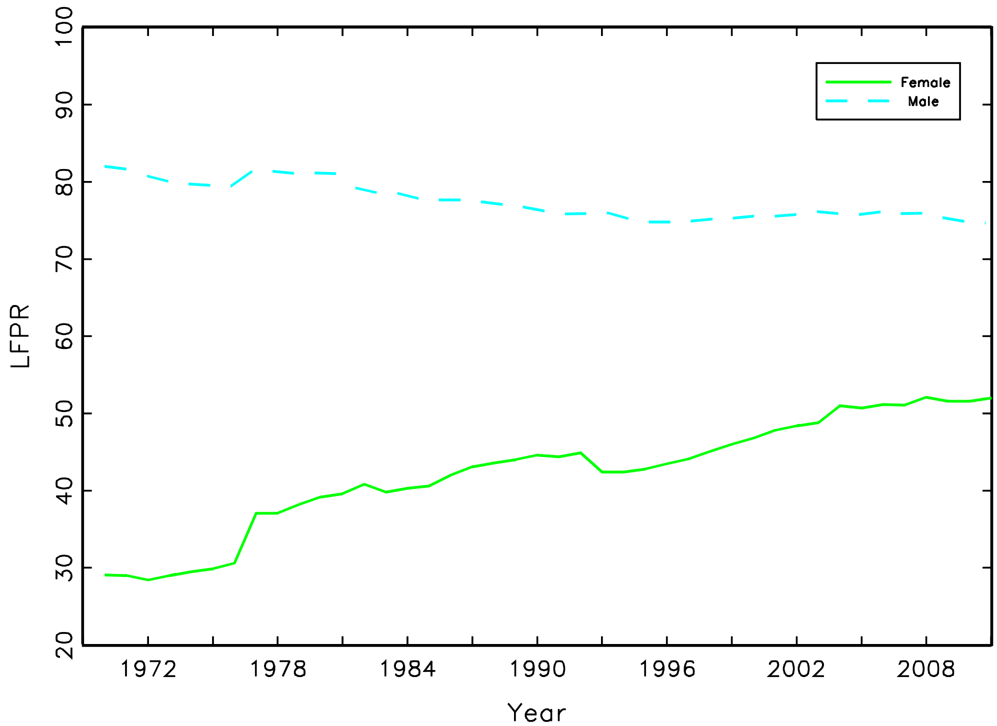

Figure 6: LFPR^M & LFPR^F for Japan, 1968–2002

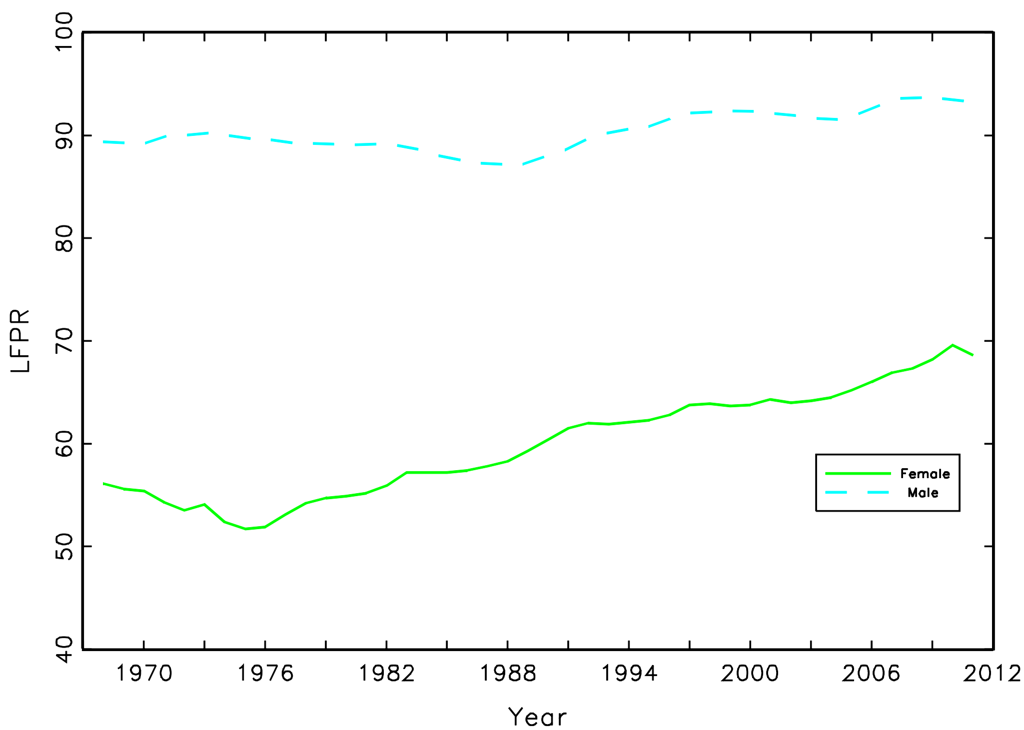

Figure 7: $LFPR^M$ & $LFPR^F$ for The Netherlands, 1971–2002

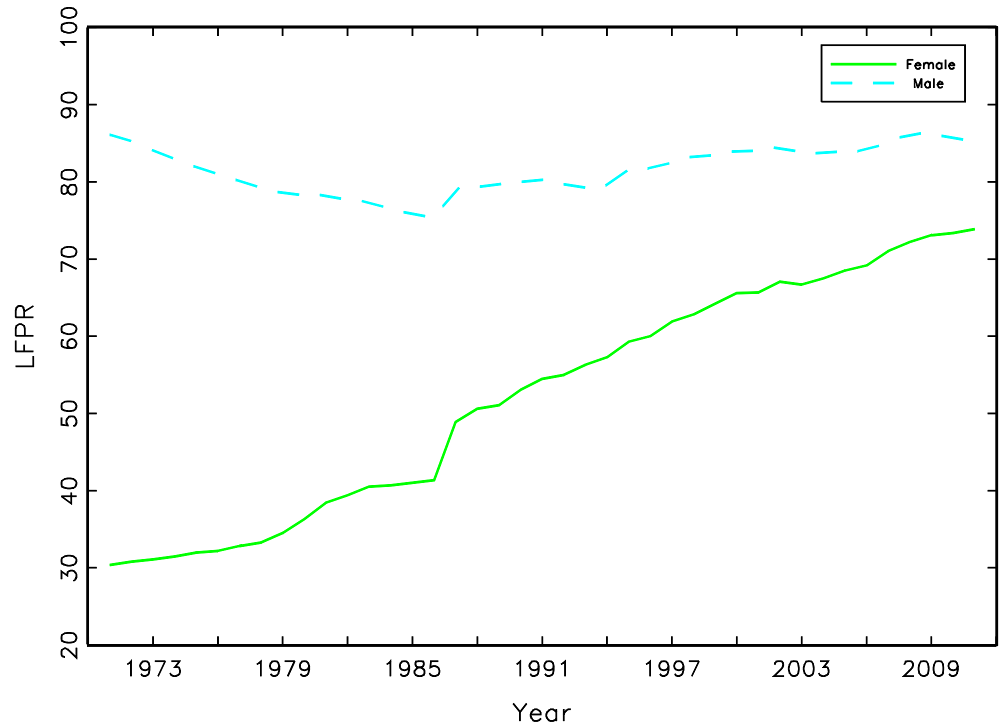

Figure 8: $LFPR^M$ & $LFPR^F$ for Norway, 1972–2002

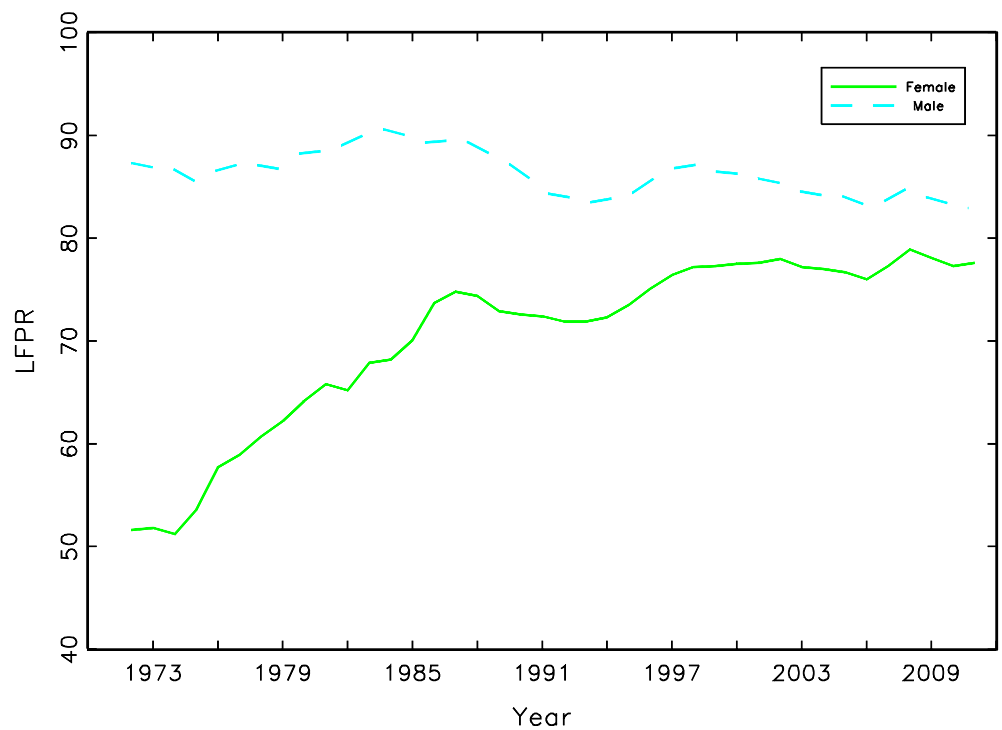

Figure 9: $LFPR^M$ & $LFPR^F$ for Portugal, 1974–2002

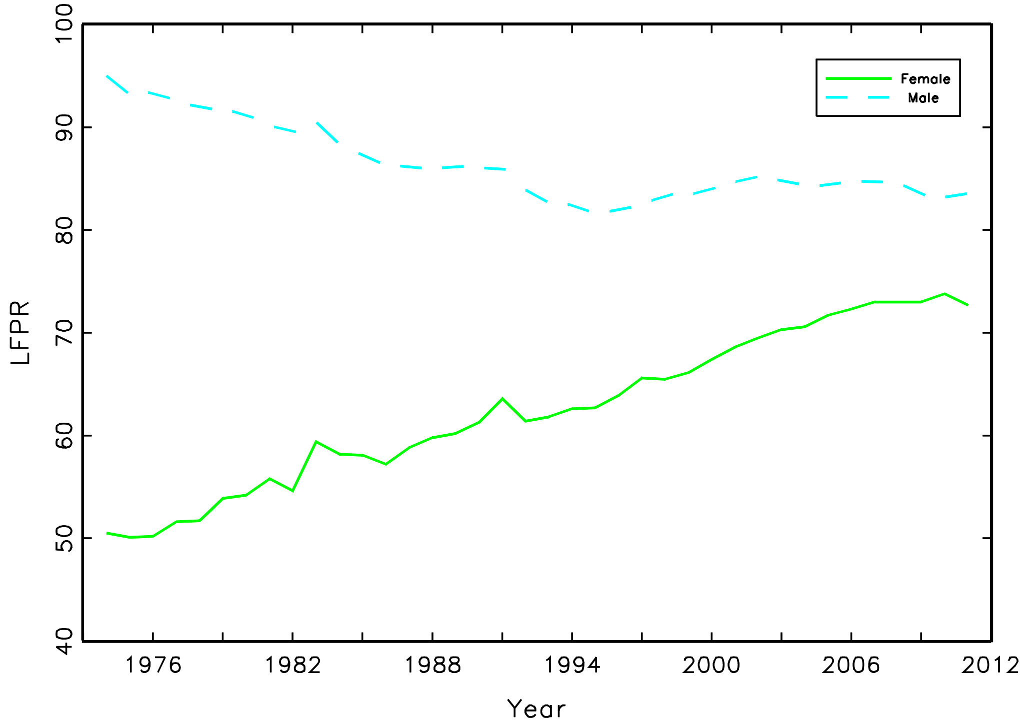

Figure 10: $LFPR^M$ & $LFPR^F$ for Spain, 1972–2002

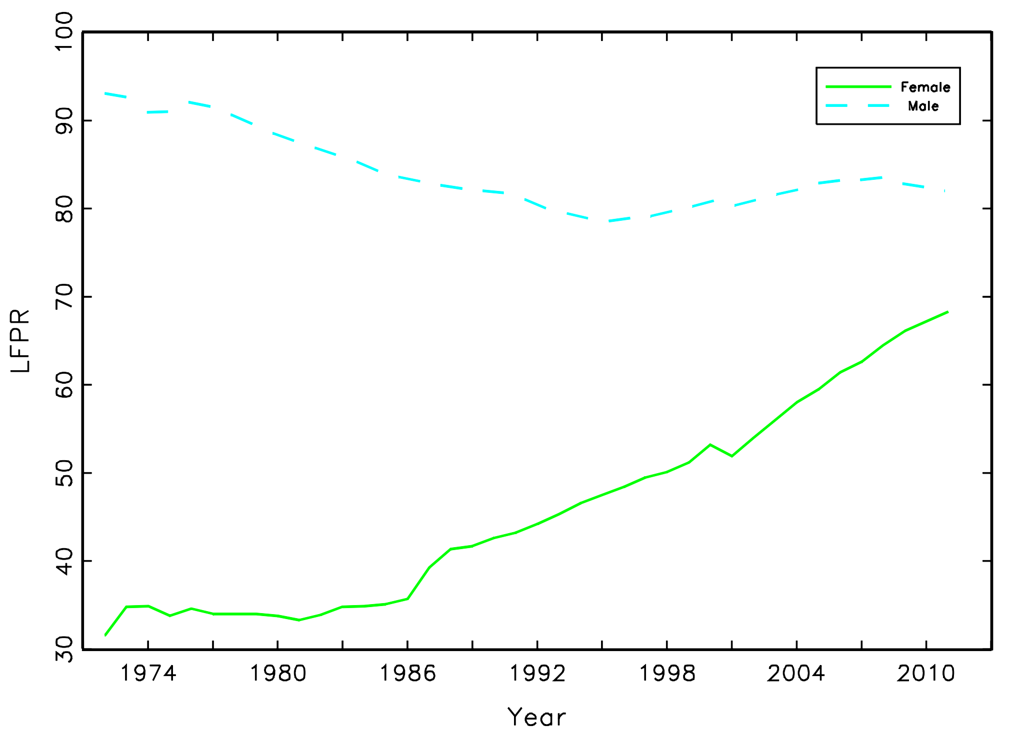

Figure 11: $LFPR^M$ & $LFPR^F$ for Sweden, 1963−2002

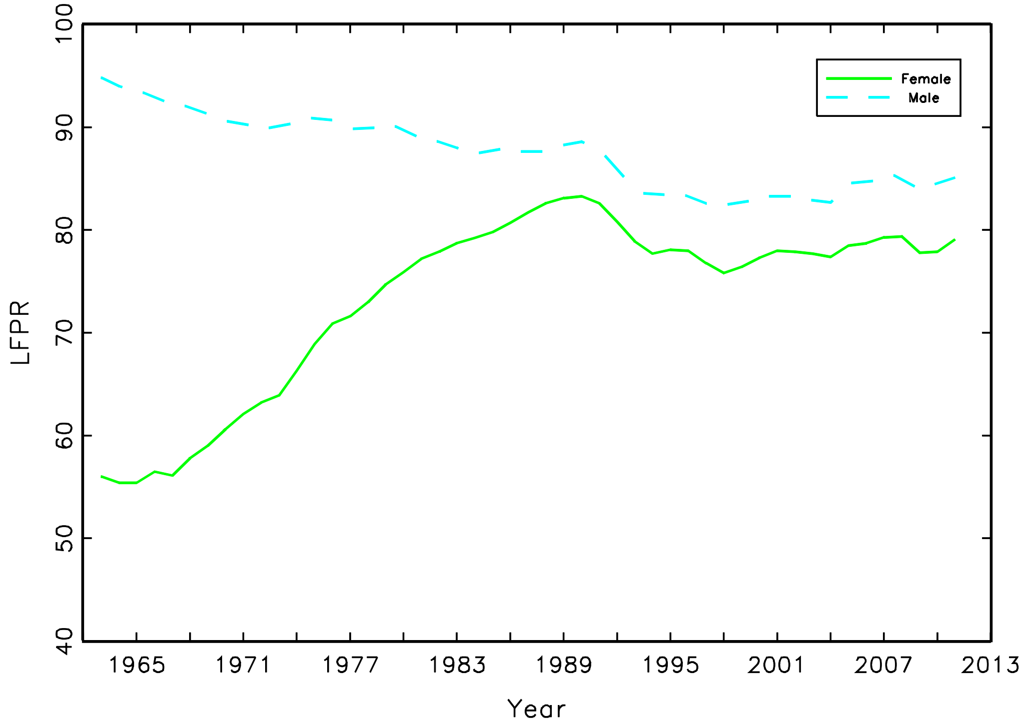

Figure 12: $LFPR^M$ & $LFPR^F$ for the U.S., 1960−2002

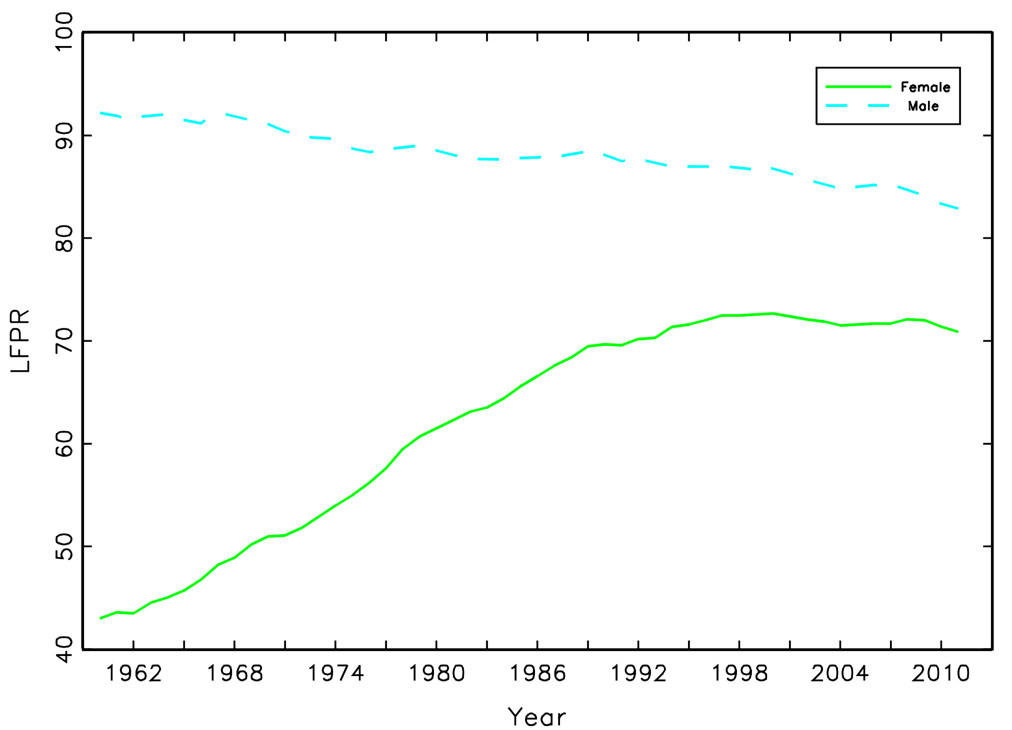

Table 2: ADF Tests for the $LFPR^M$ and $LFPR^F$ series for all Countries

		Without Trend						With Trend					
Series	T	k^*	t_μ	$\hat{\rho}$	$\hat{\mu}$	HL_ρ	k^*	t_τ	$\hat{\rho}$	$\hat{\mu}$	$\hat{\beta}$	HL_ρ
$LFPR^M$(AUS)	46	0	**-3.23[b]**	0.9352	5.48	10.34	0	-0.12	0.9947	-0.06	0.0150	130.44
$LFPR^M$(CAN)	36	0	-1.22	0.9230	6.49	8.65	0	-1.31	0.8838	9.98	-0.0079	5.61
$LFPR^M$(FIN)	49	2	-2.31	0.8673	10.42	4.87	2	-2.11	0.7278	22.32	-0.0295	2.18
$LFPR^M$(DEU)	42	0	**-4.16[a]**	0.8797	9.79	5.41	0	-1.30	0.9516	3.20	0.0295	13.97
$LFPR^M$(ITA)	42	0	-1.57	0.9251	5.63	8.90	0	-2.15	0.7899	16.75	-0.0304	2.93
$LFPR^M$(JAP)	44	1	-1.28	0.9614	3.52	17.61	1	-2.05	0.9171	7.31	0.0100	8.00
$LFPR^M$(NLD)	41	0	-1.13	0.9445	4.51	12.14	0	-3.03	0.8580	10.55	0.0491	4.53
$LFPR^M$(NOR)	40	0	-0.82	0.9437	4.74	11.96	0	-1.71	0.8538	13.06	-0.0272	4.39
$LFPR^M$(PRT)	38	0	**-2.71[d]**	0.9089	7.57	7.26	0	-1.53	0.9047	7.97	-0.0017	6.92
$LFPR^M$(ESP)	40	2	-1.50	0.9610	3.14	17.42	1	-1.16	0.9577	3.21	0.0079	16.03
$LFPR^M$(SWE)	49	0	-2.14	0.9397	5.08	11.14	1	-1.70	0.8653	12.18	-0.0223	4.79
$LFPR^M$(USA)	52	0	0.21	1.0052	-0.64	133.64	1	-2.13	0.7766	20.49	-0.0373	2.74
$LFPR^F$(AUS)	46	-	-	-	-	-	0	-1.68	0.8614	6.69	0.0877	4.64
$LFPR^F$(CAN)	36	-	-	-	-	-	1	-2.84	0.8825	7.40	0.0512	5.55
$LFPR^F$(FIN)	49	-	-	-	-	-	1	-1.86	0.9045	6.34	0.0190	6.91
$LFPR^F$(DEU)	42	-	-	-	-	-	0	-1.58	0.8730	6.18	0.0875	5.10
$LFPR^F$(ITA)	42	-	-	-	-	-	0	-1.78	0.8451	5.48	0.0753	4.12
$LFPR^F$(JAP)	44	-	-	-	-	-	1	**-3.62[c]**	0.7548	12.60	0.1061	2.46
$LFPR^F$(NLD)	41	-	-	-	-	-	0	-1.93	0.8205	5.72	0.2253	3.50
$LFPR^F$(NOR)	40	-	-	-	-	-	0	-1.20	0.9382	5.12	-0.0057	10.87
$LFPR^F$(PRT)	38	-	-	-	-	-	0	**-3.93[b]**	0.3127	34.72	0.4548	0.60
$LFPR^F$(ESP)	40	-	-	-	-	-	0	-1.04	0.9423	1.78	0.0879	11.66
$LFPR^F$(SWE)	49	-	-	-	-	-	4	-2.87	0.9367	4.91	0.0030	10.60
$LFPR^F$(USA)	52	-	-	-	-	-	1	1.14	1.0232	-0.27	-0.0329	30.22

Note: The superscripts 'a,' 'b,' 'c,' and 'd' denote respectively significance at the 1%, 2.5%, 5% and 10% significance level. The superscript * denotes near significance at the 10% level. The finite sample critical values corresponding to T=25 and T=50 were taken from Table 4.2, pp. 103 in Banerjee, Dolado, Galbraith, and Hendry (1993). The critical values for the ADF unit-root tests (t_μ) without trend: for T=25 are -2.63 at the 10% level, -3.00 at the 5% level, -3.33 at the 2.5% level, and -3.75 at the 1% level; and for T=50 are -2.60 at the 10% level, -2.93 at the 5% level, -3.22 at the 2.5% level, and -3.58 at the 1% level. The critical values for the ADF unit-root tests with trend (t_τ): for T=25 are -3.24 at the 10% level, -3.60 at the 5% level, -3.95 at the 2.5% level, and -4.38 at the 1% level; and for T=50 are -3.18 at the 10% level, -3.50 at the 5% level, -3.80 at the 2.5% level, and -4.15 at the 1% level. We extrapolated the critical values for the given sample sizes based on these critical values.

Table 3: Model M_2 Results for the $LFPR^M$ and $LFPR^F$ series of all Countries

Series	$\hat{T}_B$	$\hat{\lambda}$	k^*	$\hat{\rho}$	$t^2_{\hat{\rho}}$	F^2_T	$\hat{\sigma}$	HL_ρ
$LFPR^M$(AUS)	1998	0.85	0	0.9003	-1.42	2.33	0.3693	6.60
$LFPR^M$(CAN)	2001	0.72	1	0.6196	-3.32	5.66	0.4506	1.45
$LFPR^M$(FIN)	1988	0.53	1	0.4708	**-4.57**[d]	**10.54**[b]	0.7977	0.92
$LFPR^M$(DEU)	1990	0.50	1	0.8076	-2.83	**13.11**[a]	0.4077	3.24
$LFPR^M$(ITA)	1981	0.29	1	0.7133	-2.62	**13.38**[a]	0.5219	2.05
$LFPR^M$(JAP)	1989	0.50	1	0.7532	-3.43	**8.57**[c]	0.2845	2.45
$LFPR^M$(NLD)	1986	0.39	0	0.7310	-2.26	**25.52**[a]	0.5137	2.21
$LFPR^M$(NOR)	1990	0.48	4	0.2954	-3.94	**7.62**[c]	0.6710	0.57
$LFPR^M$(PRT)	1997	0.63	0	0.4274	-3.48	6.20	0.6485	0.81
$LFPR^M$(ESP)	2000	0.73	4	1.2050	1.67	**12.32**[a]	0.3308	3.72
$LFPR^M$(SWE)	2004	0.86	1	0.6995	-3.07	**7.07**[c]	0.5945	1.94
$LFPR^M$(USA)	1988	0.56	1	0.6262	-2.61	3.14	0.3976	1.48
$LFPR^F$(AUS)	1985	0.43	0	0.6013	-3.56	**9.59**[b]	0.4983	1.36
$LFPR^F$(CAN)	1991	0.44	4	0.5119	-2.97	4.90	0.4163	1.04
$LFPR^F$(FIN)	1990	0.57	1	0.6101	**-4.23**[*]	**7.55**[d]	0.7278	1.40
$LFPR^F$(DEU)	1990	0.50	1	0.7593	-2.99	**38.97**[a]	0.3768	2.52
$LFPR^F$(ITA)	1992	0.55	0	0.7124	-2.22	3.84	1.0945	2.04
$LFPR^F$(JAP)	1976	0.20	0	0.6610	-2.84	6.08	0.5190	1.67
$LFPR^F$(NLD)	1986	0.39	0	0.7662	-2.31	**47.59**[a]	0.5642	2.60
$LFPR^F$(NOR)	1988	0.43	3	0.4445	-3.11	6.31	0.9152	0.85
$LFPR^F$(PRT)	1982	0.24	0	0.3692	**-4.18**[d]	**13.42**[a]	0.8368	0.70
$LFPR^F$(ESP)	1986	0.38	0	0.6776	-2.68	**7.71**[d]	0.7871	1.78
$LFPR^F$(SWE)	1991	0.59	2	0.8899	-1.24	4.01	0.7267	5.94
$LFPR^F$(USA)	1992	0.63	1	0.7091	-3.45	5.76	0.3533	2.02

Note: The superscripts 'a,' 'b,' 'c,' and 'd' denote respectively significance at the 1%, 2.5%, 5% and 10% significance level. The superscript * denotes near significance at the 10% level. The finite sample critical values for $t^2_{\hat{\rho}}$ corresponding to T=50 were taken from Costantini and Sen (2013b), and the finite sample critical values for F^2_T corresponding to T=50 are taken from Costantini and Sen (2013a). We extrapolated the critical values for each unit root statistic based on the estimated break-fraction ($\hat{\lambda}$).

Table 4: Model M_0 Results for the $LFPR^M$ and $LFPR^F$ series of all Countries

Series	$\hat{T}_B$	$\hat{\lambda}$	k^*	$\hat{\rho}$	$t^0_{\hat{\rho}}$	F^0_T	$\hat{\sigma}$	HL_ρ
$LFPR^M$(AUS)	1973	0.17	0	0.8847	**-3.42**[d]	**9.07**[a]	0.3650	5.66
$LFPR^M$(CAN)	1991	0.44	0	0.7549	-1.87	3.64	0.4836	2.47
$LFPR^M$(FIN)	1988	0.53	2	0.7808	-2.78	**7.09**[b]	0.8575	2.80
$LFPR^M$(DEU)	1990	0.50	1	0.9570	-1.35	**12.70**[a]	0.4936	15.77
$LFPR^M$(ITA)	1981	0.29	1	0.8242	-2.07	**11.55**[a]	0.5531	3.59
$LFPR^M$(JAP)	2004	0.84	1	0.9293	-1.96	2.99	0.3326	9.45
$LFPR^M$(NLD)	1986	0.39	0	0.9370	-2.00	**27.99**[a]	0.5253	10.65
$LFPR^M$(NOR)	1990	0.48	4	0.6081	-2.83	**5.22**[d]	0.7880	1.39
$LFPR^M$(PRT)	1991	0.47	0	0.8886	-2.11	**8.60**[a]	0.6793	5.87
$LFPR^M$(ESP)	2000	0.73	2	0.9619	-1.51	2.56	0.5701	17.84
$LFPR^M$(SWE)	1991	0.59	0	0.8437	-2.85	**7.21**[b]	0.6299	4.08
$LFPR^M$(USA)	1974	0.29	3	1.0656	1.47	**5.49**[c]	0.3599	10.91

Note: The superscripts 'a,' 'b,' 'c,' and 'd' denote respectively significance at the 1%, 2.5%, 5% and 10% significance level. The superscript * denotes near significance at the 10% level. The finite sample critical values for $t^2_{\hat{\rho}}$ corresponding to T=50 were taken from Costantini and Sen (2013b), and the finite sample critical values for F^2_T corresponding to T=50 are taken from Costantini and Sen (2013a). We extrapolated the critical values for each unit root statistic based on the estimated break-fraction ($\hat{\lambda}$).