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INFANT MORTALITY RATES: TIME TRENDS AND FRACTIONAL INTEGRATION

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Abstract

This paper examines the time trends in infant mortality rates in a number of countries in the 20th century. Rather than imposing that the error term is a stationary I(0) process, we allow for the possibility of fractional integration and hence for a much greater degree of flexibility in the dynamic specification of the series. Indeed, once the linear trend is removed, all series appear to be I(d) with d > 0 rather than I(0), implying long-range dependence. As expected, the time trend coefficients are significantly negative, although of a different magnitude from to those obtained assuming I(0) disturbances.

Keywords: Infant mortality rates; time trends; fractional integration

JEL Classification: C22, I12

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1. Introduction

The issue of modelling trends in infant mortality rates (IMR) is still a controversial one. Obviously the IMR cannot keep declining linearly forever, since at some point it would reach the value 0 and it can never be negative. For this reason the logarithm transformation has been widely used implying an exponential decay (as in the seminal paper by Preston, 1975). The implication is a much faster decline than would be implied by a linear process. Whether or not IMRs have declined exponentially, and the statistical adequacy of a log transformation, have been examined in a recent study by Bishai and Opuni (2009). Using maximum likelihood methods, they show that only in the case of the US is the decline exponential, whilst in the other 17 countries included in their sample the best fit is obtained when IMR is linear in time. Moreover, imposing a log transform can lead to biased estimates of the relationship between IMR and GDP per capita. More recently, some papers have taken a growth regression approach to modelling IMR, finding that only primary school enrolment and vaccination rates for infants are significant factors driving it (see Younger, 2001).

However, even when imposing a log-transformation of the data the regression errors in a model with a linear trend are usually assumed to be I(0), a rather strong assumption rarely verified in empirical studies. For the purpose of the present paper, an I(0) process is defined as a covariance stationary process with a spectral density function that is positive and finite at the zero frequency. This includes standard models in time series analysis such as white noises, stationary ARs, MAs, stationary ARMAs, etc. If there is strong evidence that the series is not stationary I(0) the standard approach is then to take first differences based on the assumption that the series is I(1). However, in recent years, fractional integration or I(d) models have become plausible alternatives to the two standard (I(0) and I(1)) specifications.

In this paper we consider linear trends in the log-IMR series; however, we argue that a crucial issue in this context is the specification of the error term. In particular, instead of imposing that the error is stationary I(0) (or nonstationary I(1) in some cases) we allow for the possibility that the detrended series is I(d), where d can be any real value. This is a more general model which includes the above two as particular cases of interest.

The outline of the paper is as follows. Section 2 briefly presents the statistical model including time trends and fractional integration. Section 3 describes the data and the main empirical results, while Section 4 offers some concluding remarks.

2. Time trends and fractional integration

The standard statistical way of modelling time trends is to assume a linear function of time as in the following equation:

$$y_t = \alpha + \beta t + x_t, \quad t = 1, 2, ..., \tag{1}$$

where y_t is the observed time series (in our case, IMR), and x_t is the deviation term that is assumed to be relatively stable across time. The parameter β measures the average change in y_t per time period. In the case of the IMR series, we should expect a significantly negative value for β, which measures the average yearly reduction in the mortality rate. However, as mentioned in the introduction, in order to make valid statistical inference about β it is crucial to determine correctly the structure of the deviation term. For example, if x_t is a random variable independently drawn from a Gaussian distribution with zero mean and constant variance, the OLS estimates can be efficiently calculated, and inference is possible based on the F and t statistics (see, e.g. Hamilton, 1994, Chapter 16, and Draper and Smith, 1998). On the other hand, the detrended data may display some degree of dependence. Such behaviour can be captured by different models. One of the most widely used is the AutoRegressive process of order 1, AR(1), defined as

$$x_t \ = \ \rho x_{t-1} \ + \ \varepsilon_t , \quad t \ = \ 1, 2, ..., \qquad (2)$$

with | ρ | < 1 and white noise ε_t. This model has been widely employed in the literature because of its relation with the stochastic first-order differential equation. One can use the Prais-Wisten (1954) transformation in order to obtain a t-statistic, which converges in distribution to a N(0, 1) random variable. However, as noted by authors such as Park and Mitchell (1980) and Woodward and Gray (1993), significant size distortions appear in the test statistic when the AR coefficient in (2) is close to 1. On the other hand, if one believes that the detrended series is nonstationary, one can set ρ in (2) equal to 1, and the process is then said to be integrated of order 1 (and denoted as $x_t \sim I(1)$). Then, x_t is nonstationary, and the statistical inference should be based on its first differences, $x_t - x_{t-1}$, which are stationary. Combining now (1) and (2) (with ρ = 1) the model becomes:

$$(1 - L)\, y_t \ = \ \beta \ + \ \varepsilon_t , \quad t \ = \ 1, 2, ..., \qquad (3)$$

and one can construct another t-statistic for β.

From the comments above it is clear that it is important to determine if the detrended process x_t is stationary I(0) (even allowing for weak (ARMA)-autocorrelation) or nonstationary I(1). However, it may also be I(d) where d is a number between 0 and 1 or even above 1. This is the hypothesis examined in this study, noting that different estimates for the time trend may be obtained depending on the assumptions made about the order of integration in the detrended series.

A time series {x_t, t = 1, 2, ..., } is said to be I(d) if it can be represented as:

$$(1 \ - \ L)^d \, x_t \ = \ u_t , \qquad t \ = \ 1, 2, ..., \qquad (4)$$

and u_t is I(0). These processes (with $d > 0$) were introduced by Granger (1980, 1981), Granger and Joyeux (1980) and Hosking (1981) and since then have been widely employed to describe the behaviour of many economic time series (Diebold and Rudebusch, 1989; Sowell, 1992; Gil-Alana and Robinson, 1997; etc.).[1]

It can be showed that the polynomial on the left hand side in (1) can be expressed in terms of its Binomial expansion such that, for all real d,

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

implying that the higher is the value of d, the higher is the degree of association between observations distant in time. Thus, the parameter d plays a crucial role in determining the degree of persistence of the series. If $d = 0$ in (4), clearly $x_t = u_t$, the process is short memory, and it may be weakly (ARMA) autocorrelated with the autocorrelations decaying at an exponential rate. If d belongs to the interval (0, 0.5), x_t is still covariance stationary although the autocorrelations will take a longer time to disappear than in the previous case of I(0) behaviour; if d belongs to [0.5, 1) the process is no longer covariance stationary but it is still mean reverting in the sense that shocks will tend to disappear in the long run. Finally, if $d \geq 1$, x_t is nonstationary and not mean reverting.

Throughout this paper we estimate d in (1) and (4) using the Whittle function in the frequency domain (Dahlhaus, 1989) along with a testing Lagrange Multiplier (LM) procedure developed by Robinson (1994) that basically consists in testing the null hypothesis:

$$H_o: \; d = d_o, \qquad (5)$$

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[1] See Robinson (2003), Doukham et al. (2003) and Gil-Alana and Hualde (2009) for recent reviews of I(d) models.

in (1) and (4) for any real value d_o. This method has several advantages compared with other approaches: it tests any real value d, thus encompassing stationary ($d < 0.5$) and nonstationary ($d \geq 0.5$) hypotheses, unlike other procedures that require first differencing prior to the estimation of d. Moreover, the limit distribution is standard normal unlike most unit root methods which are based on non-standard critical values. Finally, this method is the most efficient one in the Pitman sense against local departures from the null. As in other standard large-sample testing situations, Wald and LR test statistics against fractional alternatives have the same null and limit theory as the LM test of Robinson (1994). Lobato and Velasco (2007) essentially employed such a Wald testing procedure, and, although this and other recent methods such as the one developed by Demetrescu, Kuzin and Hassler (2008) have been shown to be robust with respect to even unconditional heteroscedasticity (Kew and Harris, 2009), they require an efficient estimate of d, and therefore the LM test of Robinson (1994) seems computationally more attractive.[2]

In the following section we show that the detrended series of the log-IMR data are in fact I(d) with d statistically significantly different from zero. That means that the standard approach of estimating a linear trend using the log-transformed data and the OLS-GLS methods may lead to incorrect inferences about the time trends in the mortality rates.

3. Data and empirical results

We use data obtained from the Human Mortality Database, at the University of California, Berkeley. They are mortality rates for infants less than 1 year old, in the following countries: Australia, Austria, Belgium, Bulgaria, Canada, Czeck Republic,

[2] Other parametric estimation approaches (Sowell, 1992; Beran, 1995) were also employed for the empirical analysis producing very similar results to those obtained using the method of Robinson (1994).

Denmark, Finland, France, Hungary, Iceland, Ireland, Italy, Japan, The Netherlands, New Zeeland, Norway, Portugal, Slovakia, Spain, Sweden, Switzerland, U.K., and U.S.A..

[Insert Table 1 about here]

We report in Table 1 the sample period for each country, along with the mortality rates in two years that are far apart, namely 1950 and 2006, in order to analyse the reduction in the rates over time. It can be seen that the biggest fall occurs in the case of countries such as Slovakia, Portugal and Bulgaria that were relatively underdeveloped at the beginning of the sample and have undertaken significant reforms in the following period and experienced relatively high economic growth.

In all cases we estimate the time trends in the log-transformed data, assuming that the detrended series are I(d) where d can be any real number, thus including also integer degrees of differentiation. In other words, the specified model is:

$$\log(y_t) = \alpha + \beta t + x_t, \qquad (1 - L)^d x_t = u_t, \qquad t = 1, 2, ..., \qquad (6)$$

with white noise u_t. Although autocorrelation in u_t could also be allowed for, given the fact that the number of observations is less than 100 in most cases), autocorrelation is likely to be well described by the fractional polynomial in (6).

[Insert Table 2 about here]

Table 2 displays the estimates of d in (6) along with the 95% confidence interval of the non-rejection values of d using Robinson's (1994) approach. We disaggregate the results in male, female and total infant (<1) mortality rates. All the orders of integration are significantly greater than 0 and in many cases significantly different from 1. Therefore, the use of standard methods based on integer degrees of differentiation may produce invalid estimates of the time trend coefficients.

As for the aggregate series, there is a single country (Iceland) with a value of d below 0.5 implying stationary behaviour. In another eleven countries (New Zealand, Australia, Hungary, Switzerland, The Netherlands, Denmark, Finland, Portugal, Sweden, France and Norway) the estimated value of d is significantly below 1, implying that the unit root null hypothesis is rejected and therefore mean reversion occurs. Finally, there is another group of eleven countries (Ireland, Austria, UK, Canada, Spain, Japan, Slovakia, Bulgaria, Belgium, Italy and the US) where the unit root null hypothesis (i.e. d = 1) cannot be rejected at the 5% level, and one country, the Czech Republic, with an estimated value of d strictly above 1 (d = 1.213), the unit root null being rejected in this case in favour of d > 1.

When disaggregating the data by sex no significant differences are found, at least with respect to the degree of persistence. Specifically, for females, evidence of mean reversion (i.e., d < 1) is observed in sixteen countries (New Zealand, Iceland, Ireland, Australia, Hungar, Switzerland, Austria, Portugal, The Netherlands, Denmark, Norway, Belgium, Finland, Sweden, UK and France), and the same is found for males with the exceptions of Belgium and the UK where the unit root null cannot be rejected. It is also noteworthy that the explosive behaviour in the Czech Republic is mainly due to females since the unit root null cannot be rejected in this country for males.

[Insert Table 3 about here]

Table 3 reports for each series the estimated time trend coefficients. All of them are significantly negative. Focusing first on the aggregate series, we notice that the biggest coefficients (in absolute value) of the time trends are estimated for the Czech Republic (-0.06337) followed by Japan (-0.05855), Portugal (-0.05515), Austria (-0.05306) and Slovakia (-0.05066), while the lowest values occur for the US (-0.02922), the Netherlands (-0.02621), Norway (-0.02308), Denmark (-0.02296) and Sweden (-0.01717).

Given the different sample sizes and in order to make more meaningful comparisons between countries, in what follows we use the same sample period (1950 – 2006) for all countries. The results for the estimated values of d are presented in Table 4, while the time trends coefficients are displayed in Table 5.

[Insert Table 4 about here]

In the case of the aggregate series, the results vary substantially from one series to another. In Iceland the series may be I(0); there are nine countries with evidence of mean-reverting behaviour ($d < 1$): the Netherlands, New Zealand, Finland, Sweden, Denmark, Hungary, Austria, Australia and Portugal; the unit root null ($d = 1$) cannot be rejected in Belgium, Ireland, Switzerland, Bulgaria, Slovakia, the UK and Canada; and evidence of $d > 1$ is obtained for the Czech Republic and the US. Once more we do not find significant differences between results for males and females.

[Insert Table 5 about here]

The estimated time trend coefficients for the time period 1950 - 2006 are displayed in Table 5. For the aggregate series, the highest values (in absolute values) are those for Portugal (-0.06230), the Czech Republic (-0.06074), Spain (-0.05521), Japan (-0.05476), Italy (-0.05312) and Austria (-0.05201), while the lowest values are those for Norway (-0.02308), the US (-0.02389) and Iceland (-0.03031). Similarly, for females and males, the highest values are those for Portugal, the Czech Republic, Japan and Spain and the lowest ones those for the US and Norway.

[Insert Table 6 about here]

In Table 6 we compare the estimates of the time trends under the assumption that the detrended series are I(d) and I(0) for the aggregate data. There are substantial differences in some cases. In sixteen countries (Iceland, Japan, Sweden, Norway, Portugal, Italy, Austria, Spain, Ireland, Denmark, Switzerland, Australia, Canada, the UK, New Zealand and the US) the time trend coefficient is over-estimated when wrongly imposing the I(0) specification for the error term. On the other hand in eight countries ((Finland, Czech Republic, France, Belgium, The Netherlands, Hungary and Slovakia) the values of the time trend are under-estimated with I(0) errors. The biggest differences are found in the cases of the Czech Republic (-0.06074 with I(d) errors and -0.04120 with I(0) errors), Iceland (-0.03031 with I(d) errors and -0.04410 with I(0) ones) and Norway (-0.02308 with I(d) errors and -0.03693 with I(0) errors).

4. Conclusions

In this paper we have estimated the time trend coefficients for Infant Mortality Rates (IMR, infants less than 1 year) in 24 countries, based on the log-transformed data and using I(d) specifications of the error term. This is a general model that includes the standard cases of I(0) stationarity and I(1) nonstationarity as special cases of interest. The fact that the order of integration may be fractional allows for a greater degree of flexibility in the dynamic specification of the series. The results indicate that in all the countries examined the order of integration in the detrended series are I(d) with d strictly positive, and significantly different from zero, implying that the series display long memory behaviour. This suggests that the estimation of the time trend coefficients based on standard I(0) errors may produce invalid results because of the misspecification of the order of differentiation.

As for the orders of integration, we find that there is one country (Iceland) with an estimated value of d in the interval (0, 0.5) implying stationary behaviour. For a group of eleven countries ((New Zealand, Australia, Hungary, Switzerland, The

Netherlands, Denmark, Finland, Portugal, Sweden, France and Norway) the values of d are in the interval [0.5, 1), implying nonstationarity but mean reverting behaviour. For the remaining countries the differencing parameter is equal to or higher than 1. In general we do not observe significant differences between males and females in terms of the degree of dependence. With respect to the time trend coefficients, the most significant ones are those corresponding to the Czech Republic, Japan, Portugal, Austria and Slovakia, while the lowest values are those for the US, the Netherlands, Norway, Denmark and Sweden, more developed countries to start with and therefore with lower reductions in the mortality rates.

References

Bishai, D. and M. Opuni (2009) Are infant mortality rate declines exponential? The general pattern of 20th century infant mortality decline, Population Health Metrics, 7, 13, 1-8.

Dahlhaus. R., (1989), "Efficient parameter estimation for self-similar processes", Annals of Statistics 17, 1749-1766.

Demetrescu, M., V. Kuzin and U. Hassler (2008) Long memory testing in the time domain, Econometric Theory 24, 176-215.

Diebold, F.X., and G.D. Rudebusch (1989). "Long memory and persistence in aggregate output", Journal of Monetary Economics 24, 189-209.

Doukham, P., G. Oppenheim and M.S. Taqqu (2003), Long range dependence. Theory and Application, Boston, US.

Draper, N.R. and H. Smith (1998) Applied Regression Analysis, Third Edition, John Wiley & Sons, 706 pp.

Gil-Alana, L.A. and J. Hualde (2009) Fractional integration and cointegration. An overview with an empirical application. The Palgrave Handbook of Applied Econometrics, Volume 2. Edited by Terence C. Mills and Kerry Patterson, MacMillan Publishers, pp. 434-472.

Gil-Alana, L.A. and Robinson, P.M. (1997) Testing of unit roots and other nonstationary hypotheses in macroeconomic time series, *Journal of Econometrics* 80, 241-268.

Granger, C.W.J. (1980) "Long memory relationships and the aggregation of dynamic models." *Journal of Econometrics* 14, 227-238.

Granger, C.W.J. (1981). "Some properties of time series data and their use in econometric model specification." *Journal of Econometrics* 16, 121-130.

Granger, C.W.J. and R. Joyeux, (1980). "An introduction to long memory time series and fractionally differencing," *Journal of Time Series Analysis* 1, 15-29.

Hamilton, J.D. (1994) "Time Series Analysis", Princeton University Press.

Hosking, J.R.M. (1981). "Fractional differencing." *Biometrika* 68, 168-176.

Kew, H. and D. Harris (2009) Heteroskedasticity robust testing for a fractional unit root, Econometric Theory 25, 1734-1753.

Lobato, I. and C. Velasco (2007) Efficient Wald tests for fractional unit roots, Econometrica 75, 575-589.

Park, R.E. and B.M. Mitchell (1980) Estimating the autocorrelated error model with trended data, Journal of Econometrics 13, 185-201.

Prais, S.J. and C.B. Winsten (1954) Trend estimators and serial correlation, Cowles Commission Monograph, No. 23, New Haven CT, Yale University Press.

Preston, S.H. (1975) The changing relation between mortality and the level of economic development, Population Studies, 29, 231-248.

Robinson, P.M. (1994), "Efficient tests of nonstationary hypotheses", *Journal of the American Statistical Association*, 89, 1420-1437.

Robinson, P.M. (2003), Time series with long memory, P.M. Robinson, eds. Oxford, Oxford University Press.

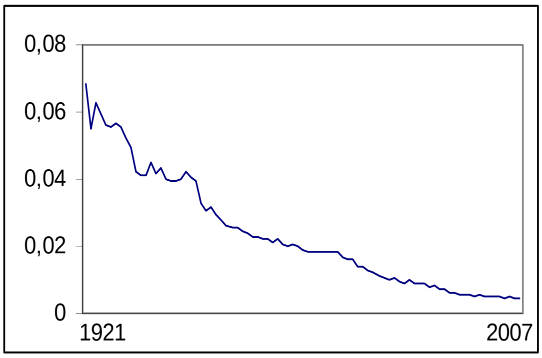
Sowell, F. (1992) Modelling long run behaviour with the fractional ARIMA model, Journal of Monetary Economics 29, 277-302.

Woodward, W.A. and H.L. Gray (1993) Distinguishing between deterministic and random trends in time series data, Computers Science and Statistics, Proceedings of the 25th Symposium on the Interface.

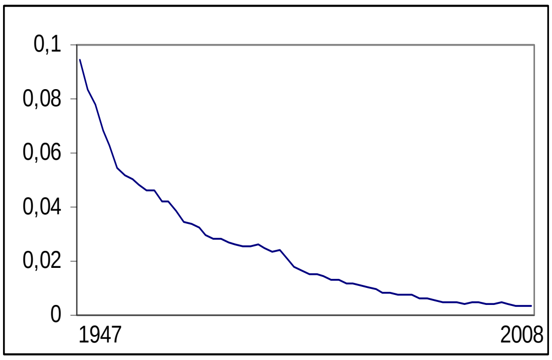
Younger, S.D. (2001) Cross-country determinants of declines in infant mortality: a growth regression approach?, Cornell University, Food and Nutrition Policy Programme, mimeo.

Figure 1: Time series plots

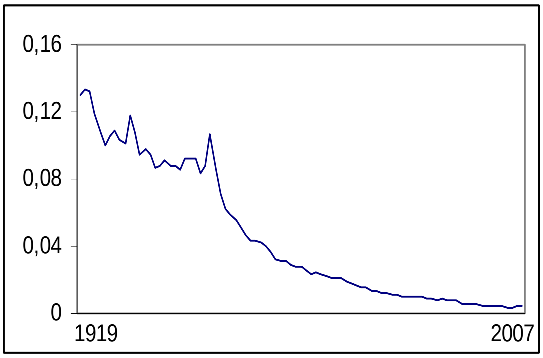
Australia



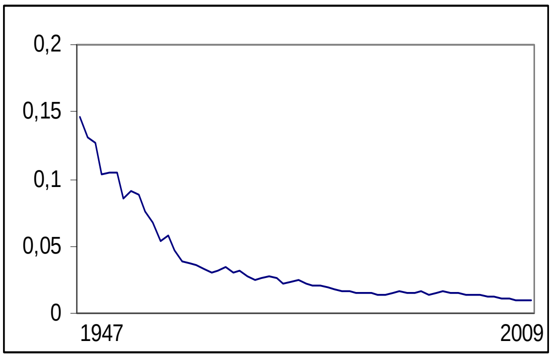
Austria



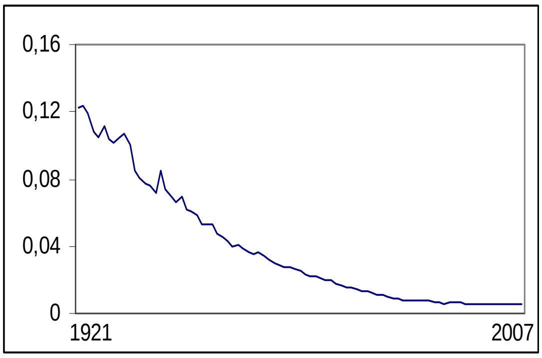
Belgium



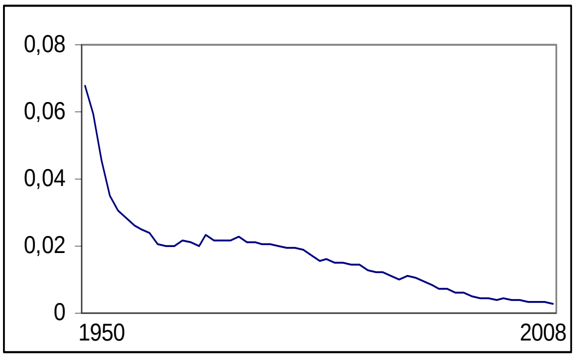
Bulgaria



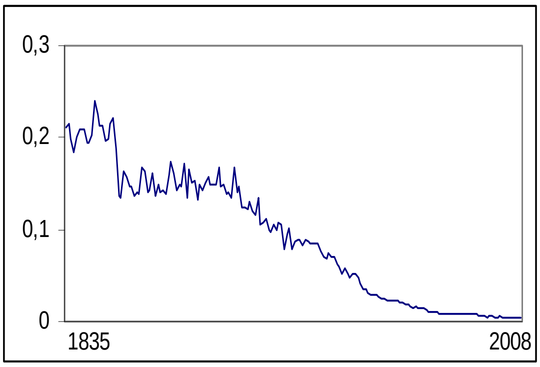
Canada



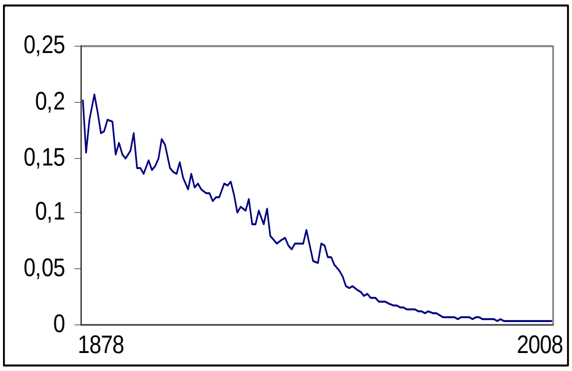
Czech Republic



Denmark



Finland



(cont.)

Figure 1: Time series plots (cont.)

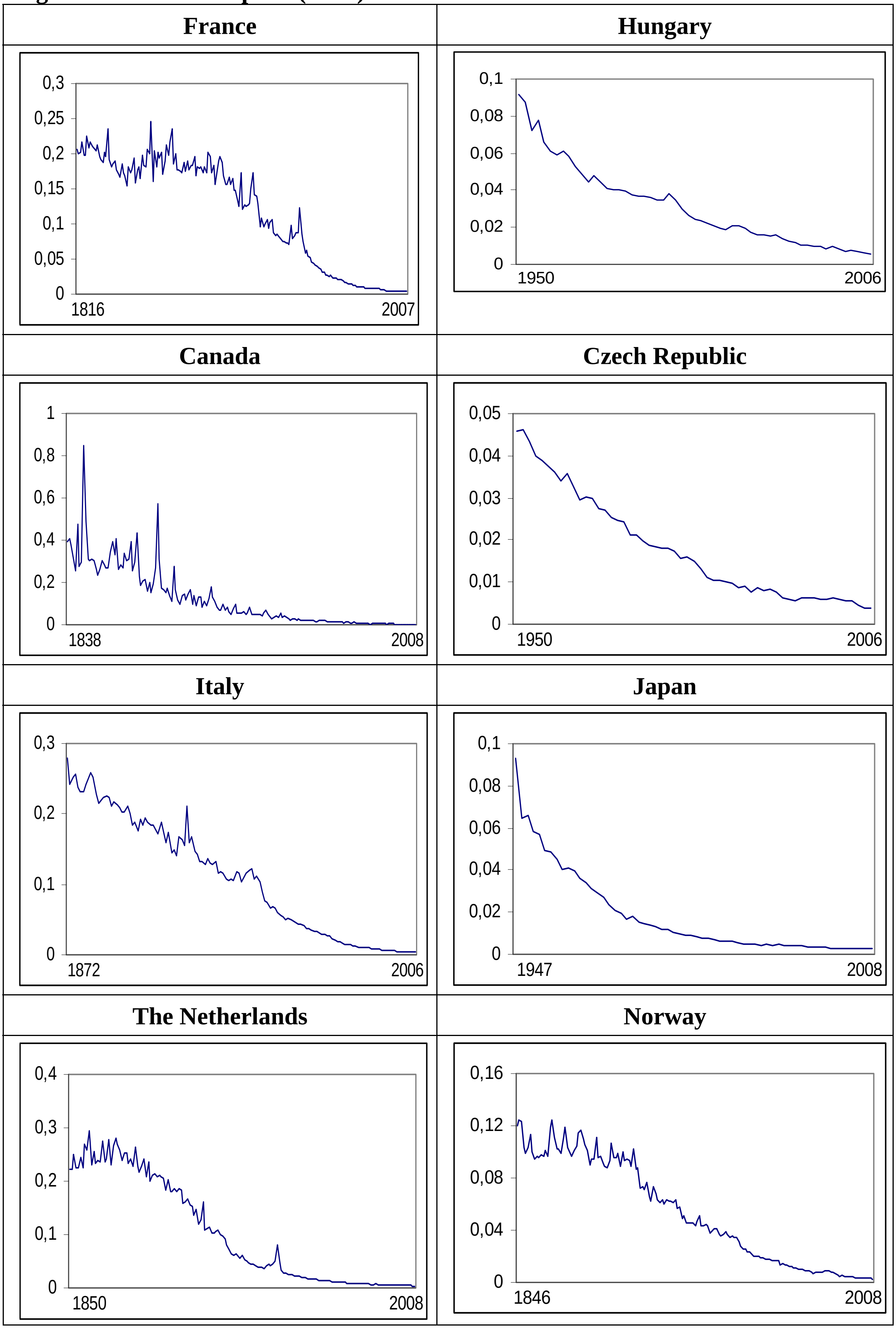


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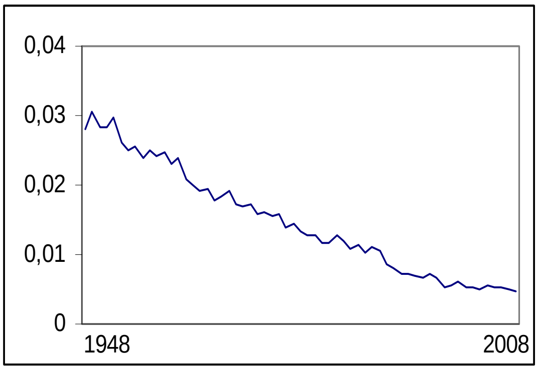
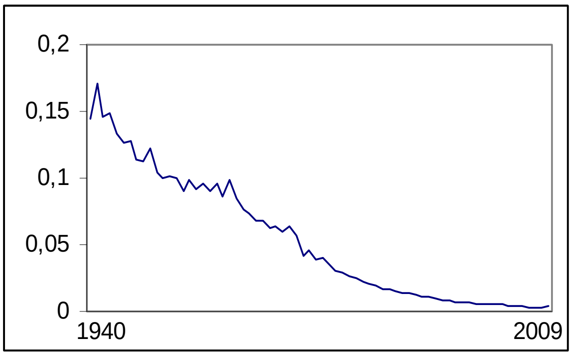
Figure 1: Time series plots (cont.)

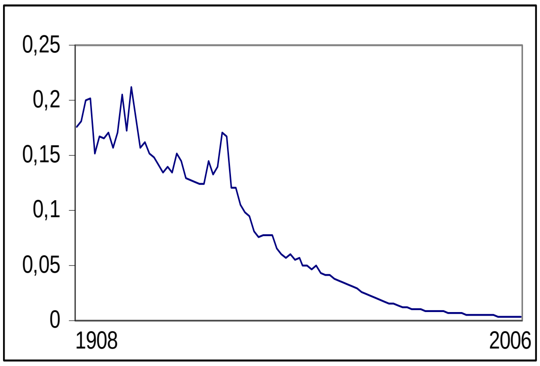
New Zealand



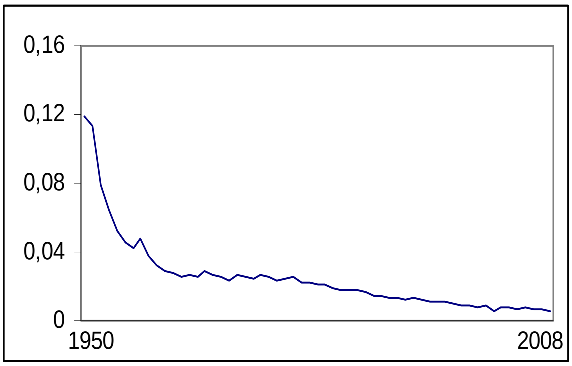
Czech Republic



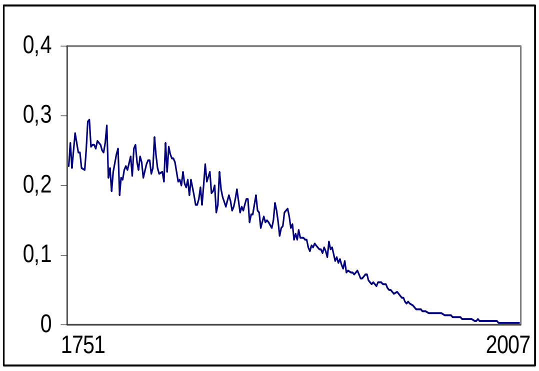
Spain



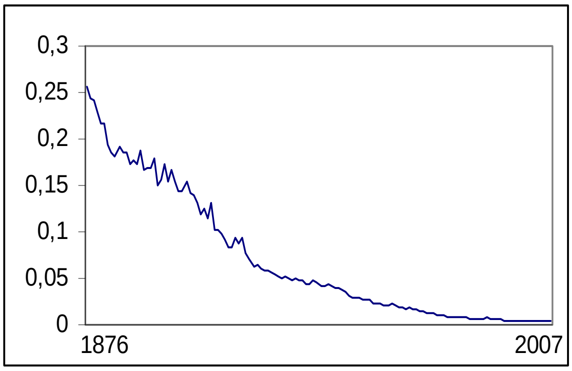
Slovakia



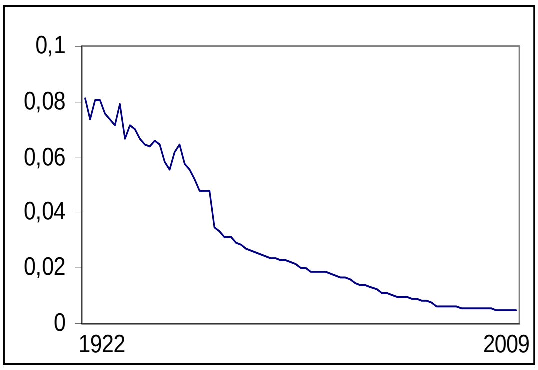
Sweden



Switzerland



UK



US

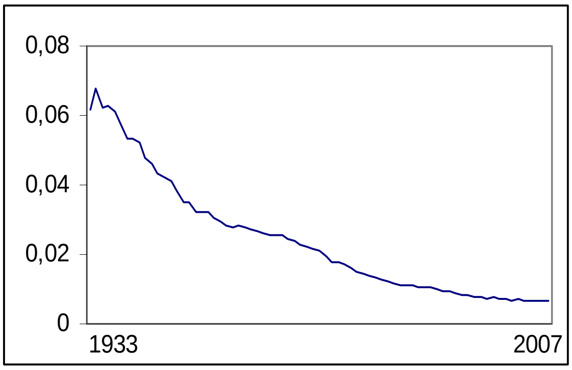


Table 1: Sample period and reduction in IMR by country

Country	Time period	1950	2006	Reduction
ICELAND	1838 - 2008	0.022790 (2)	0.001372 (1)	0.021418 (21)
JAPAN	1947 - 2008	0.058245 (15)	0.002693 (2)	0.055552 (9)
SWEDEN	1751 - 2007	0.020851 (1)	0.002857 (3)	0.017994 (24)
FINLAND	1878 - 2008	0.043277 (12)	0.002874 (4)	0.040403 (13)
NORWAY	1846 - 2008	0.025948 (5)	0.003196 (5)	0.022752 (20)
PORTUGAL	1940 - 2009	0.104441 (21)	0.003260 (6)	0.101181 (2)
CZECH REP.	1950 - 2008	0.067921 (17)	0.003380 (7)	0.064541 (6)
ITALY	1872 - 2006	0.066113 (16)	0.003468 (8)	0.062645 (8)
AUSTRIA	1947 - 2008	0.068012 (18)	0.003604 (9)	0.064408 (7)
SPAIN	1908 - 2006	0.077807 (18)	0.003681 (10)	0.074126 (5)
FRANCE	1816 - 2007	0.053602 (13)	0.003716 (11)	0.049886 (11)
IRELAND	1950 - 2006	0.045980 (13)	0.003800 (12)	0.042180 (12)
DENMARK	1835 - 2008	0.031315 (8)	0.003853 (13)	0.027462 (16)
BELGIUM	1919 - 2007	0.055066 (14)	0.004072 (14)	0.050994 (10)
NETHERLANDS	1850 - 2008	0.025320 (3)	0.004411 (15)	0.020909 (22)
SWITZERLAND	1876 - 2007	0.032827 (10)	0.004462 (16)	0.028365 (15)
AUSTRALIA	1921 – 2007	0.025426 (4)	0.004696 (17)	0.020730 (23)
CANADA	1921 - 2007	0.042973 (11)	0.005058 (18)	0.037915 (14)
UK	1922 - 2009	0.030922 (7)	0.005090 (19)	0.025832 (17)
NEW ZEALAND	1948 - 2008	0.028433 (6)	0.005172 (20)	0.023261 (19)
HUNGARY	1950 - 2006	0.091480 (19)	0.005841 (20)	0.085639 (4)
SLOVAKIA	1950 - 2008	0.118871 (22)	0.006585 (21)	0.112286 (1)
US	1933 - 2007	0.032468 (9)	0.006813 (22)	0.025655 (18)
BULGARIA	1947 - 2009	0.103598 (20)	0.010355 (23)	0.093243 (3)

Table 2: Estimates of d for each series and 95% confidence bands

Country	Female	Male	Total
ICELAND	0.410 (0.342, 0.504)	0.370 (0.314, 0.446)	0.431 (0.367, 0.519)
JAPAN	0.946 (0.854, 1.074)	0.921 (0.812, 1.075)	0.953 (0.851, 1.096)
SWEDEN	0.828 (0.791, 0.875)	0.870 (0.833, 0.918)	0.863 (0.826, 0.910)
FINLAND	0.813 (0.751, 0.901)	0.811 (0.752, 0.893)	0.850 (0.789, 0.937)
NORWAY	0.801 (0.746, 0.875)	0.866 (0.809, 0.944)	0.903 (0.842, 0.987)
PORTUGAL	0.739 (0.658, 0.848)	0.846 (0.754, 0.979)	0.862 (0.772, 0.988)
CZECH REP.	1.134 (1.004, 1.309)	1.106 (0.973, 1.282)	1.213 (1.079, 1.389)
ITALY	0.967 (0.914, 1.038)	0.964 (0.913, 1.032)	0.977 (0.922, 1.044)
AUSTRIA	0.727 (0.504, 0.998)	0.738 (0.565, 0.961)	0.816 (0.647, 1.027)
SPAIN	0.938 (0.859, 1.047)	0.932 (0.854, 1.041)	0.945 (0.866, 1.055)
FRANCE	0.853 (0.813, 0.904)	0.878 (0.837, 0.932)	0.868 (0.827, 0.920)
IRELAND	0.418 (0.284, 0.610)	0.552 (0.353, 0.833)	0.761 (0.558, 1.039)
DENMARK	0.798 (0.751, 0.861)	0.834 (0.786, 0.898)	0.847 (0.798, 0.913)
BELGIUM	0.813 (0.704, 0.971)	0.959 (0.819, 1.178)	0.972 (0.838, 1.181)
NETHERLANDS	0.788 (0.725, 0.872)	0.801 (0.738, 0.883)	0.807 (0.744, 0.891)
SWITZERLAND	0.681 (0.606, 0.781)	0.744 (0.669, 0.847)	0.768 (0.692, 0.869)
AUSTRALIA	0.578 (0.474, 0.722)	0.724 (0.613, 0.885)	0.695 (0.585, 0.852)
CANADA	0.929 (0.819, 1.084)	0.886 (0.777, 1.032)	0.929 (0.819, 1.079)
UK	0.830 (0.716, 0.978)	0.881 (0.755, 1.049)	0.881 (0.760, 1.042)
NEW ZEALAND	0.281 (0.145, 0.481)	0.627 (0.486, 0.824)	0.548 (0.412, 0.743)
HUNGARY	0.657 (0.463, 0.920)	0.727 (0.561, 0.977)	0.741 (0.560, 0.999)
SLOVAKIA	0.845 (0.703, 1.013)	0.964 (0.829, 1.136)	0.956 (0.829, 1.126)
US	1.031 (0.896, 1.214)	1.061 (0.927, 1.239)	1.059 (0.925, 1.241)
BULGARIA	0.884 (0.790, 1.012)	0.937 (0.837, 1.079)	0.956 (0.859, 1.094)

Table 3: Estimates of the time trend along with the t-values

Country	Female	Male	Total
ICELAND	-0.03125 (-16.71)	-0.03006 (-19.34)	-0.03031 (-17.44)
JAPAN	-0.05812 (-9.09)	-0.05880 (-10.58)	-0.05855 (-9.57)
SWEDEN	-0.01697 (-6.96)	-0.01711 (-6.27)	-0.01717 (-6.47)
FINLAND	-0.03391 (-8.66)	-0.03282 (-8.59)	-0.03333 (-8.07)
NORWAY	-0.02321 (-7.75)	-0.02251 (-6.68)	-0.02308 (-6.13)
PORTUGAL	-0.05673 (-12.69)	-0.05450 (-8.97)	-0.05515 (-9.30)
CZECH REP.	-0.05992 (-3.37)	-0.05696 (-3.44)	-0.06337 (-3.04)
ITALY	-0.03314 (-6.24)	-0.03220 (-6.21)	-0.03265 (-6.06)
AUSTRIA	-0.05372 (-17.03)	-0.05283 (-13.68)	-0.05306 (-13.46)
SPAIN	-0.04036 (-6.16)	-0.03963 (-6.13)	-0.03988 (-6.01)
FRANCE	-0.02114 (-5.51)	-0.02064 (-5.08)	-0.02086 (-5.27)
IRELAND	-0.04283 (-24.34)	-0.04470 (-18.81)	-0.04398 (-12.80)
DENMARK	-0.02293 (-7.68)	-0.02295 (-7.08)	-0.02296 (-6.91)
BELGIUM	-0.04061 (-10.33)	-0.03958 (-6.33)	-0.03976 (-6.41)
NETHERLANDS	-0.02633 (-8.12)	-0.02617 (-7.90)	-0.02621 (-7.78)
SWITZERLAND	-0.03181 (-19.83)	-003220 (-17.11)	-0.03194 (-16.46)
AUSTRALIA	-0.03105 (-21.84)	-0.03192 (-15.07)	-0.03152 (-16.59)
CANADA	-0.03669 (-9.08)	-0.03821 (-11.49)	-0.03735 (-9.79)
UK	-0.03319 (-10.77)	-0.03372 (-9.37)	-0.03341 (-9.39)
NEW ZEALAND	-0.03142 (-24.17)	-0.03088 (-13.75)	-0.03095 (-17.64)
HUNGARY	-0.04704 (-16.14)	-0.04767 (-14.44)	-0.04761 (-13.94)
SLOVAKIA	-0.05013 (-5.45)	-0.05009 (-4.29)	-00.0566 (-4.45)
US	-0.02920 (-7.13)	-0.02944 (-6.52)	-0.02922 (-6.60)
BULGARIA	-0.04423 (-5.60)	-0.04265 (-4.96)	-0.04350 (-4.91)

Table 4: Estimates of d and the 95% confidence interval for the time period 1950-2006

Country	Female	Male	Total
ICELAND	0.072 (-0.093, 0.311)	0.008 (-0.133, 0.201)	0.093 (-0.064, 0.352)
JAPAN	1.023 (0.939, 1.135)	0.981 (0.888, 1.108)	1.041 (0.952, 1.163)
SWEDEN	0.386 (0.215, 0.623)	0.622 (0.431, 0.905)	0.602 (0.422, 0.857)
FINLAND	0.447 (0.319, 0.636)	0.477 (0.356, 0.653)	0.580 (0.457, 0.762)
NORWAY	0.560 (0.401, 0.762)	0.749 (0.573, 0.983)	0.897 (0.723, 1.117)
PORTUGAL	0.737 (0.643, 0.859)	0.808 (0.713, 0.933)	0.843 (0.751, 0.963)
CZECH REP.	1.133 (1.003, 1.304)	1.087 (0.959, 1.255)	1.195 (1.066, 1.366)
ITALY	0.779 (0.637, 0.982)	0.808 (0.678, 0.994)	0.887 (0.745, 1.102)
AUSTRIA	0.618 (0.409, 1.020)	0.675 (0.503, 0.913)	0.745 (0.568, 0.998)
SPAIN	0.799 (0.679, 0.963)	0.807 (0.689, 0.963)	0.845 (0.728, 1.002)
FRANCE	0.849 (0.708, 1.046)	0.857 (0.676, 1.137)	0.906 (0.732, 1.161)
IRELAND	0.418 (0.284, 0.610)	0.552 (0.353, 0.833)	0.761 (0.558, 1.039)
DENMARK	0.524 (0.399, 0.8702)	0.562 (0.415, 0.765)	0.709 (0.557, 0.930)
BELGIUM	0.303 (0.044, 0.641)	0.623 (0.396, 0.958)	0.680 (0.405, 1.084)
NETHERLANDS	0.435 (0.313, 0.604)	0.413 (0.289, 0.576)	0.510 (0.387, 0.678)
SWITZERLAND	0.596 (0.460, 0.793)	0.684 (0.530, 0.901)	0.821 (0.664, 1.046)
AUSTRALIA	0.651 (0.539, 0.795)	0.759 (0.644, 0.910)	0.756 (0.644, 0.903)
CANADA	1.027 (0.909, 1.205)	1.070 (0.946, 1.260)	1.121 (0.997, 1.322)
UK	0.891 (0.751, 1.083)	1.058 (0.891, 1.301)	1.107 (0.942, 1.344)
NEW ZEALAND	0.281 (0.142, 0.484)	0.626 (0.467, 0.858)	0.538 (0.398, 0.745)
HUNGARY	0.657 (0.463, 0.920)	0.727 (0.561, 0.977)	0.741 (0.560, 0.999)
SLOVAKIA	0.842 (0.697, 1.016)	0.964 (0.827, 1.142)	0.959 (0.824, 1.131)
US	1.279 (1.107, 1.558)	1.327 (1.178, 1.555)	1.356 (1.192, 1.629)
BULGARIA	0.870 (0.769, 1.019)	0.959 (0.838, 1.134)	0.952 (0.843, 1.111)

Table 5: Estimates of the time trend along with the t-values for the time period 1950-2006

Country	Female	Male	Total
ICELAND	-0.03125 (-16.71)	-0.03006 (-19.34)	-0.03031 (-17.44)
JAPAN	-0.05480 (-8.60)	-0.05499 (-10.07)	-0.05476 (-8.83)
SWEDEN	-0.03609 (-22.77)	-0.03722 (-15.84)	-0.03655 (-17.75)
FINLAND	-0.04718 (-29.72)	-0.04774 (-22.11)	-0.04784 (-20.79)
NORWAY	-0.02321 (-7.75)	-0.02251 (-6.68)	-0.02308 (-6.13)
PORTUGAL	-0.06336 (-12.68)	-0.06142 (-10.47)	-0.06230 (-10.44)
CZECH REP.	-0.05969 (-3.27)	-0.05457 (-3.46)	-0.06074 (-3.05)
ITALY	-0.05391 (-21.10)	-0.05289 (-16.89)	-0.05312 (-15.28)
AUSTRIA	-0.05140 (-22.21)	-0.05242 (-16.02)	-0.05201 (-16.33)
SPAIN	-0.05542 (-18.28)	-0.05479 (-15.50)	-0.05521 (-15.54)
FRANCE	-0.04777 (-6.73)	-0.04811 (-14.40)	-0.04784 (-13.64)
IRELAND	-0.04283 (-24.34)	-0.04470 (-18.81)	-0.04398 (-12.80)
DENMARK	-0.03641 (-14.54)	-0.03816 (-16.28)	-0.03633 (-8.11)
BELGIUM	-0.04672 (-47.53)	-0.04760 (-24.19)	-0.04718 (-23.11)
NETHERLANDS	-0.03097 (-24.04)	-0.03171 (-33.10)	-0.03145 (-26.80)
SWITZERLAND	-0.03682 (-14.40)	-0.03757 (-13.01)	-0.03630 (-9.49)
AUSTRALIA	-0.03122 (-13.40)	-0.03114 (-9.74)	-0.03106 (-10.39)
CANADA	-0.03719 (-6.04)	-0.03853 (-6.19)	-0.03765 (-5.35)
UK	-0.03199 (-9.50)	-0.03255 (-6.13)	-0.03149 (-5.42)
NEW ZEALAND	-0.03191 (-22.86)	-0.03125 (-12.91)	-0.03160 (-17.39)
HUNGARY	-0.04704 (-16.14)	-0.04767 (-14.44)	-0.04761 (-13.94)
SLOVAKIA	-0.05019 (-5.29)	-0.05044 (-4.17)	-0.05094 (-4.28)
US	-0.02451 (-3.14)	-0.02483 (-2.88)	-0.02389 (-2.62)
BULGARIA	-0.03971 (-4.93)	-0.04221 (-4.35)	-0.04090 (-4.43)

Table 6: Comparisons of the time trends with I(d) and I(0) errors

Country	I(d) errors	I(0) errors
ICELAND	-0.03031 (-17.44)	-0.04410 (-35.671)
JAPAN	-0.05476 (-8.83)	-0.05665 (-45.453)
SWEDEN	-0.03655 (-17.75)	-0.03734 (-61.240)
FINLAND	-0.04784 (-20.79)	-0.04683 (-58.562)
NORWAY	-0.02308 (-6.13)	-0.03693 (-41.299)
PORTUGAL	-0.06230 (-10.44)	-0.06674 (-46.333)
CZECH REP.	-0.06074 (-3.05)	-0.04120 (-22.958)
ITALY	-0.05312 (-15.28)	-0.05579 (-87.880)
AUSTRIA	-0.05201 (-16.33)	-0.05246 (-74.349)
SPAIN	-0.05521 (-15.54)	-0.05842 (-78.515)
FRANCE	-0.04784 (-13.64)	-0.04781 (-92.968)
IRELAND	-0.04398 (-12.80)	-0.04417 (-63.834)
DENMARK	-0.03633 (-8.11)	-0.03764 (-47.218)
BELGIUM	-0.04718 (-23.11)	-0.04671 (-103.571)
NETHERLANDS	-0.03145 (-26.80)	-0.03102 (-69.381)
SWITZERLAND	-0.03630 (-9.49)	-0.03906 (-53.003)
AUSTRALIA	-0.03106 (-10.39)	-0.03383 (-42.943)
CANADA	-0.03765 (-5.35)	-0.04313 (-49.501)
UK	-0.03149 (-5.42)	-0.03557 (-59.553)
NEW ZEALAND	-0.03160 (-17.39)	-0.03261 (-48.550)
HUNGARY	-0.04761 (-13.94)	-0.04509 (-59.855)
SLOVAKIA	-0.05094 (-4.28)	-0.03983 (-25.694)
US	-0.02389 (-2.62)	-0.03215 (-57.956)
BULGARIA	-0.04090 (-4.43)	-0.03666 (-20.385)