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JOÃO TOVAR JALLES*

Is There A Stable Long-run Relationship Between Unemployment And Productivity?**

Abstract

This paper assesses whether productivity and unemployment have a stable long-run relationship. We explore a panel of 19 OECD countries between 1970 and 2012 and rely on recently developed time series econometric methods. Our findings suggest that unemployment and productivity are non-stationary in levels and in many individual cases these series are cointegrated, even after accounting for possible structural breaks. For many individual countries the long-run effect seems to be generally positive. There is also evidence of two-way causality, but the stronger directional relationship runs from unemployment to productivity.

Keywords: *stationarity, structural breaks, cointegration, DOLS, Granger causality*

1. Introduction

Productivity, in its broadest meaning, refers to an economy's ability to efficiently convert inputs into outputs. Macroeconomists devote a lot of their attention to productivity-related variables in order to date productivity slowdowns and revivals as well as to account for their causes and consequences. The empirical literature dealing with productivity distinguishes between the 1948-1973 period - the

* Ph.D., Center for Globalization and Governance, Nova School of Business and Economics, Portugal

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Golden Age -, and the post-1973 period - characterized by a productivity slowdown. The most common explanation for such a slowdown is based on the oil price shocks (Griliches 1988; Fisher 1988; Dolmas et al. 1999). There are, however, other explanations for the post-1973 productivity slowdown,[1] and the current paper is particularly interested in those related to labour market conditions, such as the increase in female labour force participation (Bowman, 1991) and the increase in the growth rate of labour inputs (Romer 1987).

Our main goal is to evaluate whether (labour) productivity and unemployment have a stable long-run relationship. Despite the existence of several theoretical papers relating these two variables (see Section 2), the empirical evidence remains small and/or inconclusive. To this end we use a panel of 19 advanced countries between 1970 and 2012. We rely on recent time series techniques, such as (individual) unit root and cointegration tests allowing for structural breaks, Granger-causality and Dynamic OLS estimation.

Empirical findings suggest that unemployment and labour productivity are non-stationary in levels and in many individual cases these series are cointegrated, even after accounting for possible structural breaks. Long-run cointegration estimates seem to suggest a positive co-movement between unemployment and productivity, therefore providing evidence in support of those models (Caballero and Hammour, 1994) which suggest a positive (long-run) co-movement between these two variables. Causality is found to be bi-directional in many countries, with the stronger relationship running from unemployment to productivity.

In Section 2 we review the literature, and in Section 3 outline the econometric methodology. Section 4 presents and discusses the empirical results, and the final section offers conclusions.

2. Literature Review

In terms of theoretical contributions, a recent paper by Barnichon (2010) shows that, by means of a New-Keynesian search model of unemployment with nominal rigidities and variable labour effort, technology shocks can generate a positive unemployment-productivity correlation, whereas non-technology shocks tend to produce the opposite. Moreover, the author argues that the correlation between unemployment and productivity changed in the mid-1980s

[1] We can refer here to the growth of the underground economy and under-reporting of income (Fichtembaum, 1989); demand constraints (Walker and Vatter, 1989); under-measurement of output in the services sector (Griliches, 1994); price mis-measurement (Gordon, 1996); and a decrease inf energy consumption (Beaudreau, 1998).

from significantly negative to significantly positive.[2] Despite the existence of a variety of factors that are likely to influence this relationship (e.g. interest rates, hiring and firing costs, income taxation, non-labour costs, unemployment benefits, saving behaviour), one can distinguish between two opposite views on whether periods of economic expansion lead to higher productivity in the long-run. The first is that during times of low economic activity we have smaller productivity (King and Rebelo 1988 and Stiglitz 1993).[3] On the other hand, the New-Schumpeterian approach does not support the view that unemployment is negatively correlated with output (Caballero and Hammour 1994).

Empirically, the strict focus on the correlation between these two series has led to mixed results. Earlier studies (for the US economy or for a small set of advanced countries), based on the neo-Marxian hypothesis that average labour productivity is significantly related to labour market conditions, is attributed to Weisskopt et al. (1983) and Weisskopt (1987). Taking a broader view, Bean and Pissarides (1993) examined cross-country correlations for the OECD economies between unemployment and labour productivity for the period 1955-1985. There was no clear correlation except over the period 1975-85, where a weak negative coefficient appears to be significant. However, such cross-sectional analyses are fragile in nature since country-specific effects can weaken the underlying relations (due to different institutional and economic factors which are unrelated to productivity). Looking at time series data for a particular country seems more reasonable, especially if we take into account the relative constancy of institutions within each nation over time. Caballero (1993) looked at quarterly time series evidence from the US and UK between 1966 and 1989. The author used a Hodrick-Prescott filter to remove the high-frequency components, however the evidence he found was not conclusive. For medium frequencies, both countries demonstrated a positive relation between the two variables under scrutiny.[4] More recently, Brauninger and Pannenberg (2002) took a generalised augmented Solow-type model and found that unemployment reduces long-run productivity. They then confirmed this theoretical result empirically with a panel of 13 OECD countries between 1960 and 1990. Muscatelli and Tirelli (2001) applied Structural Time Series Models to 11 OECD countries between 1955 and 1990 and found evidence in favour of those theories predicting a negative co-movement between unemployment and productivity.

[2] Other studies include the pioneering work by Gali (1999), followed by more recent papers from Holly and Petrella (2008) and Gali and Gambetti (2009).

[3] Stadler's (1990) learning-by-doing model emphasizes the link between employment and growing productivity through human capital investments.

[4] Other approaches have used VAR models, but these ended up having mixed results as well (Saint-Paul, 1997).

3. Methodology

3.1 Unit Roots and Structural Breaks

When it comes to stationarity assessments, in addition to standard Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) unit root tests - for purposes of robustness and completeness[5] - we also conduct the four tests (M-tests) proposed by Ng and Perron (2001) (NP) based on modified information criteria (MIC): the modified Phillips-Perron test MZ_{α}; the modified Sargan-Bhargava test (MSB); the modified point optimal test MP_T; and the modified Phillips-Perron MZ_T. These improve the PP-tests both with regard to size distortions and power.

We then resort to unit root tests allowing for breaks and we begin with the Zivot-Andrews (1992) (ZA) test. This endogenous structural break test is a sequential test which utilizes the full sample and uses a different dummy variable for each possible break date. The break date is selected where the t-statistic from the ADF test of unit root is at a minimum (most negative). Consequently a break date will be chosen where the evidence is least favourable for the unit root null.[6] We complement this with the modified ADF test proposed by Vogelsang and Perron (1998) (VP), also allowing for one endogenously determined break. Finally, we take the two-break unit root test described by Clemente, Montanes and Reyes (1998) (CMR). This tests the null of unit root against the break-stationary alternative hypothesis and provides us supplementary insights vis-a-vis the conventional unit root tests, which do not account for any break in the data.

For the unit root tests that allow for one or two endogenously determined breaks it is assumed that the shift can be modelled by a dummy variable $DU_t = 0$ for $t \leq TB$ and for $t > TB$, where TB is the shift date (time break). In the time series literature, two generating mechanisms of shifts are distinguished - the additive

[5] This test is especially appropriate under certain dynamic data structures and when their random components are not white noise.

[6] The critical values in Zivot and Andrews (1992) are different from the critical values in Perron (1989): the selection of the time of the break is treated as the outcome of the estimation procedure, rather than predetermined exogenously.

outlier (AO) and innovational outlier (IO) models. The former results in an abrupt shift in the level, whereas the latter allows for a smooth shift from the initial level to a new level. Although both results are reported, we will mainly discuss tests constructed for AO models.[7]

However, it is important to recognize some important drawbacks in both earlier unit root tests, particularly, the ZA and VP tests. In particular, with respect to the VP test it has been shown that the critical values are substantially smaller in the *I(0)* case than in the *I(1)* case, therefore suggesting that the test is conservative in the *I(0)* case. The solution was then to devise a procedure that would have the same limit distribution in both cases. This was first attempted by Vogelsang (2001), but simulations provided support for the lack of power in the *I(1)* case. Perron and Yabu (2009) (PY) were more successful in this endeavour by proposing a new test for structural changes in the trend function of the time series without any prior knowledge of whether the noise component was stationary or integrated. This newer test has better properties in terms of size and power.[8]

3.2. Cointegration, Stability and Causality

Consider the following (cointegrating-relationship) regression:

$$prod_{it} = \alpha_i + \beta u_{it} + \varepsilon_{it} \,. \qquad (1)$$

where $prod_{it}$ is the log of productivity and u_{it} the log of unemployment. ε_{it} is a standard iid disturbance term.

Given the nonstationarity of each individual time series (to be tested and confirmed in Section 4), the relevant question becomes whether a linear combination of these variables is stationary. If such a combination exists, productivity and unemployment become cointegrated, which implies that the variables are attracted to a stable long-run (equilibrium) relation and any deviation from this relation reflects short-run (temporary) disequilibria.

We test for cointegrating (long-run) relations between productivity and unemployment using the Johansen and Juselius (1990) methodology. This

[7] As discussed in Vogelsang and Perron (1998), the AO framework may be preferable to the IO statistics.

[8] We thank Pierre Perron and Tomoyoshi Yabu for providing their GAUSS code.

approach estimates the long-run attracting set in a VAR context, that incorporates both the short- and long-run dynamics of the various models. However, and as in the case of unit roots, a test for co-integration that does not take into account possible breaks in the long-run relationship will have lower power. The test will tend to under-reject the null of no co-integration if there is a co-integration relationship that has changed at some time during the sample period. Therefore, in order to further evaluate the previous results one should also entertain the possibility that the series are co-integrated, but that the linear combination has shifted at an unknown point in the data sample; in other words, that there might be a relevant break date. Following Gregory and Hansen (1996), the hypothesis of a structural shift in the co-integration relationships was then studied.[9] In order to estimate the parameter β in (1) we resort to the method of Dynamic Ordinary Least Squares (DOLS) of Stock and Watson (1993), following the methodology proposed by Shin (1994).[10]

As has been emphasized by Bruggemann et al. (2003), it is important to formally investigate the stability of the cointegrating vectors further once a long-run relationship has been identified. The temporal stability of estimated relations is also indicative of the usefulness of these estimated relations for policy (forecasting) purposes. Hansen and Johansen (1993) outline a procedure that formally tests the constancy of cointegrating vectors in the context of Full Information Maximum Likelihood (FIML) estimations. Holding the short-run dynamics of the model constant, the procedure then treats these estimates as the null hypothesis in consecutive recursive tests. In this way, any rejection of the null of cointegration stability (constancy) should emanate from a breakdown in the long-run relation, rather than from any positive shift in the underlying short-run dynamics (Hoffmann et al., 1995). We apply this approach to test the stability of the cointegrating relation.

By taking a VAR approach we can further extract two important additional tools: Granger-causality tests and Variance Decompositions. Many tests of Granger-type causality have been derived and implemented to test the direction of causality – Granger (1969). These tests are grounded in asymptotic theory.[11] Also, it is well documented that the exclusion of relevant variables induces spurious significances and inefficient estimates. In dealing with these problems, and for robustness purposes, we employ the Toda and Yamamoto (1995) and Dolato and Lutkepohl (1996) approach for Granger causality. They

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[9] We thank Bruce Hansen for making the GAUSS routine available.

[10] This method has the advantage of providing a robust correction to the possible presence of endogeneity in the explanatory variable, as well as of serial correlation in the error terms of the OLS estimation.

[11] For further discussions, see Toda and Phillips (1994).

suggest a technique that is applicable irrespective of the integration and cointegration properties of the system. The method involves using a Modified Wald statistic for testing the significance of the parameters of a VAR(s) model (where *s* is the lag length in the system).[12]

We follow Rambaldi and Doran (1996) in formulating these tests. Defining $d_{\max}$ as the maximum order of integration in the system, a VAR($k + d_{\max}$) has to be estimated to use the Modified Wald test for linear restrictions on the parameters of a VAR(*k*) which has an asymptotic χ^2 distribution.[13] In our case, we will run a 2 variables' VAR, with *k*=2 (AIC-based) and $d_{\max} = 1$, but for the sake of notation simplicity we denote them as $y_i, i = 1,2$. For our VAR(3) we estimate the following system of equations:

$$\begin{bmatrix} y_{1t} \\ y_{2t} \end{bmatrix} = A_0 + A_1 \begin{bmatrix} y_{1t-1} \\ y_{2t-1} \end{bmatrix} + A_2 \begin{bmatrix} y_{1t-2} \\ y_{2t-2} \end{bmatrix} + A_3 \begin{bmatrix} y_{1t-3} \\ y_{2t-3} \end{bmatrix} + \begin{bmatrix} e_{y_1} \\ e_{y_2} \end{bmatrix}$$

The above system of equations is estimated via the seemingly unrelated regression (SUR) method. This test consists of taking the first *k* VAR coefficient matrix (but not all lagged coefficients) to make Granger causal inference. If, for example, we want to test that y_{2t} does not Granger-cause y_{1t}, the null hypothesis will be $H_0 : a^{(1)}{}_{12} = a^{(2)}{}_{12} = 0$, where $a^{(i)}{}_{12}$ are the coefficients of $y_{2t-i}, i = 1,2$.

[12] As demonstrated by Toda and Yamamoto (1995), if variables are integrated of order *d*, the usual selection procedure is valid whenever $k \geq d$. Thus, if $d = 1$, the lag selection is always consistent.

[13] The traditional F tests and its Wald test counterpart to determine whether some parameter of a stable VAR model are jointly zero are not valid for non-stationary processes, as the test statistics do not have a standard distribution (Toda and Phillips, 1994).

4. Empirical Results

First, our data for a set of 19 advanced economies comes from the OECD Stat. The two main variables of interest are unemployment and (labour) productivity, measured as output per worker (both in logs).

Starting with an analysis of stationarity properties, Table 1 presents the results for several individual unit root tests allowing for none, one or two structural breaks in the underlying series. In general, unemployment series are I(1) in levels, with the exception of Belgium and Switzerland for the ADF test and Portugal for the PP test. Sweden and the US are the only two countries for which the null of stationarity is rejected in the case of the NP test(s). If one turns to tests allowing for breaks, then depending on the test we may get different results, with the overwhelmingly conclusion that most series keep their I(1) status (with the exception of Spain and Sweden), and don't reject the null of break stationarity for the ZA, VP and CMR tests. One can also note the different power attributed to the PY2009 test (particularly as the ZA and VP are conservative in the I(0) case and show a lack of power in the I(1) case), where in all but three cases we reject the null of unit root. Turning to the labour productivity series we find similar results, with the non-rejection of the null of unit root in levels for most countries (with the exception of Portugal and Spain). We observe fewer rejections of the null of unit root in the break-type tests (Portugal and Switzerland for the ZA test).

Table 1. Unit Root Tests and Structural Breaks 1970-2012

Series	ADF		PP		NP				ZA	VP(AO)	VP(IO)	CMR(AO)	CMR(IO)	PY2009
	Levels	FD	Levels	FD	MZa	MZt	MSB	MPT						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
Unemployment														
Australia	-1.44	-1.35	-1.97	-6.83***	-1.70	-0.68	0.40	35.93	1992	1976	1973	1976, 1984	1973, 1980	1981***
Austria	-1.25	-2.39	-1.38	-4.04**	-8.28	-1.93	0.23	11.31	1982	1984	1979**	1983, 1992	1979, 1989**	1980***
Belgium	-4.15**	-3.12	-1.96	-3.12	-7.68	-1.95	0.25	11.87	1988	1991	1986	1984, 1991	1986, 1995	1996
Canada	-1.74	-4.45***	-1.03	-4.79***	-7.54	-1.80	0.23	12.35	1982	1979**	1973	1979, 1992	1973, 1980	1980***
Denmark	-2.09	-4.79***	-2.37	-4.79***	-6.39	-1.77	0.27	14.24	1994	1996**	1992	1988, 1995	1985, 1992	1993***
Finland	-2.78	-5.91***	-1.56	-2.81	-5.53	-1.47	0.26	15.96	1991	1994	1989**	1979, 1993**	1974, 1989**	1990
France	-1.10	-4.55***	-0.99	-4.43***	-2.92	-1.02	0.35	26.32	1981	1978**	1973**	1976, 1982	1973, 2000**	1983***
Greece	-3.17	-3.64**	-3.19	-3.83**	-10.13	-2.07	0.20	9.75	1998	1993	1990**	1992, 1997	1990, 1994	1995***
Ireland	-0.56	-4.10**	-0.59	-3.67**	-3.23	-1.07	0.33	24.20	1991	1984	1978	1984, 2002	1972, 1978	1992***
Italy	0.49	-6.15***	1.02	-6.33***	0.31	0.14	0.47	57.48	1986	2005	1973	1980, 2002	1973, 2000	1988***
Japan	-2.12	-4.06**	-1.77	-4.05**	-14.65*	-2.55	0.17	7.08	1998	1995	1991	1977, 1995	1973, 1991	1997***
Netherlands	-2.26	-5.03***	-2.22	-3.23*	-8.52	-1.97	0.23	10.98	1988	1977	1978	1984, 2000	1979, 1986	1984***

Norway	-0.88	-4.70***	-1.06	-3.82**	-5.43	-1.39	0.25	16.01	1996	1984	1979	1980, 1990	1979, 1986	1994***
Portugal	0.90	-2.76	-4.59***	-5.47***	-0.27	-0.36	1.31	314.95	1987	1977	1976	1989, 2004	1976, 1985	1982***
Spain	-1.31	-3.21*	-1.11	-3.30*	-4.49	-1.38	0.30	19.39	1984**	1979	1973**	1977, 1982	1973, 1995**	1982***
Sweden	-2.99	-3.74**	-1.77	-3.21*	-24.78***	-3.44***	0.13***	4.10**	1992**	1994**	1989**	1983, 1993	1979, 1989**	1990***
Switzerland	-5.39***	-5.21***	-1.54	-4.55***	-9.38	-2.08	0.22	10.05	1979	1977	1972	1976, 1992**	1973, 1990	1979
UK	-1.74	-6.41***	-1.20	-3.69**	-2.31	-0.96	0.41	34.43	1980	2002	1973	1977, 1998	1973, 1995	1982***
US	-3.04	-4.43***	-2.48	-5.68***	-14.34*	-2.64*	0.18*	6.52*	1986	1976	1972	1976, 1996	1972, 1981	1982***
Labour Productivity														
Australia	0.55	-6.41***	0.81	-6.40***	2.20	2.53	1.14	110.76	1990**	2000	1991	1980, 1995	1976, 1991	1988***
Austria	-1.99	-5.81***	-1.97	-5.80***	1.61	1.74	1.08	90.01	1981	1996	1994	1985, 1997	1986, 1996	1979***
Belgium	-0.94	-5.80***	-0.94	-5.80***	1.68	2.19	1.31	129.81	1988	1991	1986	1985, 1997	1985, 1995	1979*
Canada	0.53	-4.08***	0.42	-3.98***	1.51	0.94	0.62	33.79	1982	1996	1992	1986, 1996	1986, 1996	1980**
Denmark	0.69	-5.63***	0.68	-5.63***	2.11	2.20	1.03	91.03	1994	1995	1992	1983, 1995	1982, 1992	1992***
Finland	-0.19	-3.68***	0.12	-3.08**	0.77	0.38	0.49	21.67	1991	1990	1992	1983, 1998	1981, 1995	1990
France	-1.33	-3.64**	-2.11	-4.81***	1.01	0.79	0.78	45.38	1988	1991	1984	1985, 1998	1984, 1996	1991***
Greece	-0.92	-4.70***	-1.17	-4.63***	2.37	1.33	0.56	32.44	1993	2001	1998	1972, 2000	1995, 1999	1979***

Ireland	-0.29	-4.16***	0.05	-4.02***	1.00	0.63	0.63	31.54	1985	1998	1993	1991, 1998	1986, 1993	1981
Italy	-2.27	-6.54***	-1.90	-6.94***	-7.36	-1.74	0.23	12.66	1994	1989	1986	1981, 1992	1983, 1991	1993*
Japan	-2.35	-4.34***	-2.07	-4.23***	-5.56	-1.56	0.28	16.09	1988	1985	1983	1980, 1989	1983, 2002	1991***
Netherlands	-1.91	-4.72***	-1.91	-4.58***	-11.30	-2.36	0.20	8.12	1981	1996	1995	1993, 1998	1986, 1995	1979***
Norway	-2.56	-6.44***	-2.28	-6.82***	-8.94	-2.10	0.23	10.21	1994	1993	1989	1981, 1994	1982, 1991	1992*
Portugal	-4.12**	-4.26***	-2.66	-6.06***	-34.55	-4.09	0.11	2.95	1991**	1989	1984	1981, 1993	1984, 1993	1986
Spain	-3.53*	-3.50*	-2.79	-3.61**	-13.06	-2.55	0.19	6.98	1981	1998	1993	1986, 1998	1985, 1996	1979***
Sweden	-1.44	-4.54***	-1.44	-4.47***	-2.05	-0.84	0.41	35.27	1980	2001	1992	1988, 2001	1982, 1996	1992***
Switzerland	-2.42	-5.40***	-2.49	-5.61***	-7.66	-1.93	0.25	11.93	1991***	2001	1996	1978, 1999	1977, 1996	1989***
UK	-1.73	-4.73***	-1.21	-5.14***	-7.12	-1.79	0.25	12.91	1980	1995	1991	1989, 1999	1981, 1992	1979***
US	-2.43	-4.64***	-1.52	-4.93***	-9.41	-2.14	0.22	9.79	1982	1995	1981	1985, 1995	1981, 1992	1980

Note: All variables are in logs. ADF critical values: -4028, -3.445, -3.145 for 1, 5 and 10% levels respectively. For the Ng-Perron test (NP), none of the test statistics are significant at the usual levels. The critical values are taken from Ng and Perron (2001), Table 1 and the autoregressive truncation lag (zero) has been selected using the modified AIC. The ZA test statistic reported is the minimum Dickey-Fuller statistic calculated across all possible breaks in the series, when both a break in the intercept and the time trend is allowed for. The year in parenthesis denotes the year when this minimum DF statistic is obtained. The 1% critical value is -5.57 and the 5% critical value is -5.08. As for the VP test, "AO" means addictive outlier and "IO" means innovational outlier and critical values are taken from Perron and Vogelsang (1992), in particular, -3.56 (AO) and -4.27 (IO) for 5% level. As for CMR the 5% critical value is -5.49 (both AO and IO), also taken from Perron and Vogelsang (1992). In column 10 we run the Perron-Yabu (PY) unit root test. For the structural-break type tests only dates are presented and when applicable, a statistically significant symbol is added. The null in the non-break type tests is of unit root. The null in the break-type tests is of unit root against the break stationary alternative hypothesis.

Source: author's calculations.

Having covered stationarity, we move to cointegration issues by analysing the relationship between unemployment and productivity. Table 2 presents the results for the Johansen-Juselius cointegration test. We find evidence of one cointegrating relationship in six countries (Austria, France, Ireland, Italy, Japan and Spain). Moreover, in these cases the results from the Hansen-stability test did not reject the null hypothesis that the series are cointegrated at conventional levels (with p-values larger than 20%).

Table 2. Johansen-Juselius Cointegration Tests: Productivity and Unemployment

	labprod									
Null	Alternative	Aus	Aut	Bel	Can	Den	Fin	Fra	Gre	Ire
$r=0$	$r\geq 1$	27.03*	24.04	19.94	20.43	21.16	15.93	37.88*	22.72	27.13*
$r\leq 1$	$r\geq 2$	3.11	7.83	6.54	5.93	7.07	5.27	11.64	7.06	6.03
$\lambda_{\max}$										
$r=0$	$r=1$	23.91*	16.21	13.39	14.50	14.08	10.65	26.23*	15.11	21.09*
$r\leq 1$	$r=2$	3.11	7.83	6.54	5.93	7.07	5.27	11.64	7.06	6.03
Cointegration*		Yes	No	No	No	No	No	Yes	No	Yes

(Cont.)

	labprod										
Null	Alternative	Ita	Jap	Net	Nor	Por	Spa	Swe	Swi	UK	US
$r=0$	$r\geq 1$	18.96*	36.46*	20.79	15.26	18.96	16.43*	19.63	21.37	18.00	19.67
$r\leq 1$	$r\geq 2$	0.02	5.34	5.60	6.62	3.97	0.54	8.53	7.73	6.78	7.19
$\lambda_{\max}$											
$r=0$	$r=1$	18.94*	31.11*	15.18	8.64	14.98	15.89*	11.10	13.64	19.38	12.47
$r\leq 1$	$r=2$	0.02	5.34	5.60	6.62	3.97	0.54	8.53	7.73	12.51	7.19
Cointegration*		Yes	Yes	No	No	No	Yes	No	No	No	No

Note: * denotes rejection of the null hypothesis at the 5% level (based on MacKinnon-Haug-Michelis p-values).

Source: author's calculations.

As previously discussed, we further test the hypothesis of a structural shift in the cointegration relationship for all countries in our sample by using the Gregory and Hansen (1996) procedure. Table 3 presents our results. After taking into account the possibility of breaks in the series, we get rejections of the null of no cointegration in eight countries for the ADF* statistic.

Table 3. Testing for regime shifts in cointegration of productivity and unemployment: Gregory-Hansen

	Labour Productivity			
Country	**ADF test**		**Phillips Test**	
	ADF^* stat	Estimated break date	Z_α^* stat	Estimated break date
Australia	-4.34	1986	-27.23	1987
Austria	-4.04	1975	-24.13	1974
Belgium	-5.24	1994	-25.90	1994
Canada	-4.56*	1996	-28.56	1997
Denmark	-6.82***	1992	-32.28	1992
Finland	-4.60*	1995	-28.12	1995
France	-4.14	1982	-20.14	1974
Greece	-3.89	1985	-22.08	1987
Ireland	-5.15**	1974	-23.08	1974
Italy	-4.26	1981	-25.23	1981
Japan	-4.00	1995	-19.20	1996
Netherlands	-4.25	1983	-20.44	1981
Norway	-4.94**	1993	-30.50	1992
Portugal	-4.22	1999	-15.87	1989
Spain	-4.65*	1983	-21.96	1974
Sweden	-4.05	1977	-24.38	1977
Switzerland	-4.97**	1989	-29.94	1989
UK	-5.39**	1983	-27.20	1983
US	-4.16	1998	-25.35	1999

Note: ADF^* and Z_α^* refer to the Augmented Dickey-Fuller (ADF) and to the Phillips Z_α^* tests statistics; null of no cointegration. *, ** and *** denote significance at 10, 5 and 1% levels, respectively, using the critical values from Gregory and Hansen (1996), table 1.

Source: author's calculations.

We are now in a position to estimate the parameter β in Eq. (1). The estimation is made using the DOLS of Stock and Watson (1993) as previously described. The results of the estimation of this equation for each country, in terms of the coefficient β and the statistic C_μ, a LM statistic from the DOLS residuals which tests for deterministic cointegration (i.e., when no trend is present in the regression), appear in Table 4. Two main results can be obtained from the Table. First, since all the cointegration statistics are highly significant at usual levels, the null of deterministic cointegration is rejected. And, second, the estimates of β are, in 9 out of 11 cases, positive. Up to this point our results provide evidence in support of those models (Caballero and Hammour, 1994) which suggest a positive (long-run) co-movement between productivity and unemployment. That is, this favours New-Schumpeterian theories that suggest that prolonged recessions, which are typically associated with high unemployment, foster long-run productivity improvements.

Table 4. Estimation of long-run relationships between productivity and unemployment: Stock-Watson-Shin cointegration

Country	Labour productivity		
	β	$\bar{R}^2$	C_μ
Australia	0.21 (0.03)***	0.78	4.29 (0.07)***
Austria	0.09 (0.08)	0.53	4.48 (0.24)***
Belgium	-0.61 (0.38)	0.14	6.35 (0.99)***
Canada	0.16 (0.12)	0.29	4.21 (0.37)***
Denmark	-0.53 (0.09)***	0.76	5.90 (0.20)***
Finland	0.22 (0.05)***	0.63	4.13 (0.13)***
France	0.20 (0.06)***	0.70	4.06 (0.22)***
Greece	0.31 (0.11)**	0.51	3.86 (0.31)***
Ireland	0.20 (0.15)	0.37	4.21 (0.34)***
Italy	0.37 (0.09)***	0.75	3.48 (0.32)***
Japan	0.34 (0.05)***	0.79	3.54 (0.19)***
Netherlands	-0.04 (0.04)	0.16	4.85 (0.10)***
Norway	0.24 (0.05)***	0.67	4.37 (0.10)***
Portugal	-0.18 (0.32)	0.06	4.94 (0.80)***
Spain	0.09 (0.05)	0.43	4.37 (0.17)***
Sweden	0.16 (0.03)***	0.66	4.30 (0.07)***
Switzerland	0.02 (0.01)***	0.56	4.72 (0.01)***
United Kingdom	-0.15 (0.09)*	0.48	5.15 (0.29)***
United States	0.10 (0.36)	0.09	4.39 (1.39)***

Note: The C_μ is the Shin (1994) LM statistic, which tests for deterministic cointegration. The critical values are taken from Shin (1994), Table 1, for m=1. Standard errors are in parentheses, adjusted for long-run variance. The long-run variance of the cointegrating regression residuals was estimated using the Barlett window with $l = 5 \approx INT(T^{1/2})$ as proposed by Newey and West (1987). The number of leads and lags selected was $q = 3 \approx INT(T^{1/3})$ as proposed in Stock and Watson (1993). *, ** and *** denote significance at 10, 5 and 1% levels, respectively.

Source: author's calculations.

Our final exercise is to explore the causality direction between our measures of productivity and unemployment. Tables 5.a and 5.b present our results for both the standard Granger causality test and also the Toda-Yamamoto test. In general, the evidence suggests stronger effects running from unemployment to productivity, but in some countries a two-way causality is found (e.g. Australia, Canada, Finland, UK and US in Tables 5.a and 5.b).

Table 5.a Granger causality tests

Country\Dep. Var.	**Labour productivity**			
	$u \rightarrow prod$	**Yes/No**	$prod \rightarrow u$	**Yes/No**
Australia	15.28***	Yes	24.19***	Yes
Austria	3.63	No	5.48*	Yes
Belgium	10.23***	Yes	9.42***	Yes
Canada	10.50***	Yes	11.08***	Yes
Denmark	9.86***	Yes	6.66*	Yes
Finland	10.30***	Yes	28.24***	Yes
France	1.46	No	4.43	No
Greece	9.88***	Yes	0.42	No
Ireland	0.72	No	4.48	No
Italy	2.92	No	11.10***	Yes
Japan	3.28	No	31.59	No
Netherlands	0.01	No	1.50	No
Norway	11.38***	Yes	1.31	No
Portugal	2.28	No	0.67	No
Spain	0.39	No	1.80	No
Sweden	9.16**	Yes	6.29**	Yes
Switzerland	5.81*	Yes	4.30	No
United Kingdom	9.85***	Yes	22.78***	Yes
United States	16.77***	Yes	12.00***	Yes

Note: In these tests the null is of non-Granger causality. These tests are based on a VAR with lag equal to 2, as identified using different lag-length criteria. *, ** and *** denote significance at 10, 5 and 1% levels, respectively.

Source: author's calculations.

Table 5.b Toda–Yamamoto causality tests

Country\Dep. Var.	Labour productivity			
	$u \to prod$	**Yes/No**	$prod \to u$	**Yes/No**
Australia	41.65***	Yes	30.41***	Yes
Austria	0.40	No	1.81	No
Belgium	2.93	No	10.53***	Yes
Canada	10.13***	Yes	13.74***	Yes
Denmark	1.32	No	1.54	No
Finland	6.16**	Yes	10.69***	Yes
France	1.83	No	5.37*	Yes
Greece	1.53	No	0.11	No
Ireland	0.47	No	4.98*	Yes
Italy	3.25	No	4.89*	Yes
Japan	2.23	No	29.16***	Yes
Netherlands	1.60	No	6.56**	Yes
Norway	7.95**	Yes	0.69	No
Portugal	1.54	No	4.59	No
Spain	0.02	No	1.69	No
Sweden	4.02	No	3.95	No
Switzerland	2.51	No	0.09	No
United Kingdom	11.62***	Yes	7.86**	Yes
United States	12.56***	Yes	5.76*	Yes

Note: In these tests the null is of non-Granger causality. These tests are based on a VAR(3) – see the main text for details. *, ** and *** denote significance at 10, 5 and 1% levels, respectively.

Source: author's calculations.

5. Conclusions

This paper has empirically uncovered the existence of a stable long-run relationship between productivity and unemployment in several economies within a set of 19 OECD countries between 1970 and 2012. By applying recently developed time series econometric methods, empirical findings reveal that unemployment and labour productivity are non-stationary in levels (but stationary in first-differences, hence I(1)) and in many individual cases unemployment and productivity series are cointegrated, even after accounting for possible structural breaks. Long-run cointegration estimates seem to suggest a positive co-movement between the levels of unemployment and productivity. Hence, our results provide evidence in support of those models which suggest a positive (long-run) co-movement between productivity and unemployment. Even though causality is found to be bi-directional in many cases, the stronger relationship runs from unemployment to productivity.

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Streszczenie

CZY ISTNIEJE STABILNY DŁUGOOKRESOWY ZWIĄZEK MIĘDZY BEZROBOCIEM A PRODUKTYWNOŚCIĄ?

Artykuł jest próbą ustalenia czy istnieje stabilny długookresowy związek między produktywnością a bezrobociem, Badania obejmują dane dotyczące 19 państw OECD, pochodzące z lat 1970-2012 i są oparte o najnowsze ekonometryczne metody analizy szeregów czasowych. Wyniki badań wskazują, że poziomy bezrobocia i produktywności cechują się niestacjonarnością a w licznych indywidualnych przypadkach szeregi te są skointegrowane, nawet po uwzględnieniu możliwych załamań strukturalnych. W przypadku wielu indywidualnych państw efekty długoterminowe wydają się być generalnie pozytywne. Istnieją również dowody występowania przyczynowości dwukierunkowej, ale silniejszy ukierunkowany związek zachodzi między bezrobociem a produktywnością.

***Słowa kluczowe**: stacjonarność, załamania strukturalne, kointegracja, DOLS, przyczynowość w sensie Grangera*