A panel data approach to price-value correlations

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Abstract

Resorting to stationary and non-stationary panel data econometrics we offer tests for "Ricardo's 93% theory of value" for 10 OECD countries over different time ranges. The theory does not find empirical support.

Keywords: value, price, fixed effects model, random effects model, panel unit root tests, panel cointegration tests.

JEL Codes: C21, C43, C52, D46.

1 Introduction

Measures of price-value correlations have recently been the subject of keen debate. Recent reviews of the literature have been offered by Kliman (2002, 2004) and Tsoulfidis and Paitaridis (2009). These contributions also highlight underlying theoretical issues, which are of paramount importance for Marxian/classical economics. We focus here on empirical studies only.

By way of introduction, we may usefully refer to Shaikh (1984) and Cockshott and Cottrell (1997), who estimated the following model

$$\ln P_j = \alpha + \beta \ln D_j + \epsilon_j \tag{1}$$

where j is a sectoral index, P are aggregate prices – measured by "gross output" series - D are aggregate monetary values series (whose measure is illustrated below), α and β are constants, and ϵ_j is an error term. If sectoral values are the main determinants of sectoral prices, it will follow that: (i) $\alpha = 0$; (ii) $\beta = 1$; (iii) the R^2 of (1) is large. The set of these three predictions has been termed "Ricardo's 93% Theory of Value".

Earlier contributions (Cockshott and Cottrell, 1997, 1998; Petrovic, 1987; Tsoulfidis and Maniatis, 2002), based on cross-sectional regressions and input-output data, found strong empirical support for the above predictions.

However, this evidence has more recently been questioned by Kliman (2002) who, on the basis of national accounts data, argued that prices and values tend to be higher in larger sectors and lower in smaller ones. In other words, industry size drives the strong connection between prices and values. When industry total costs are used to deflate prices and values, the support for the three predictions above vanishes.

Diaz and Osuna (2005-6, 2007) interpreted the role of size in the correlation between industry prices and values in a different way. Consider the following equation

$$\ln \frac{p_i}{p_j} = \alpha + \beta \ln \frac{d_i}{d_j} + u_i \tag{2}$$

where p_j and p_i are unit prices[1] of the jth and ith commodities respectively, d_j and d_i are unit values, u is a stochastic error, and α and β are parameters. The jth commodity is the numeraire, which is common to all the observations, so that the error is not indexed by j. In order to estimate equation (1), one has to manipulate (2) so as to obtain

$$\ln \frac{p_i q_i}{p_j q_j} = \alpha + \beta \ln \frac{d_i q_i}{d_j q_j} + (1-\beta) \ln \frac{q_i}{q_j} + u_i \tag{3}$$

where q_j and q_i are the physical quantities of output by industries j and i respectively. However, q_j and q_i cannot be observed, and their value may vary according to measurement units. Therefore estimates of (1) based on (3) would be affected by an omitted variable problem. Under these circumstances, different attempts to solve this problem by deflating industrial prices and values would lead to different results and therefore to indeterminacy. In

[1] Diaz and Osuna (2007, 391) referred to production prices. However, in section 4 they used Kliman's data where market prices are regressed on values.

other words, the conclusions by Kliman (2002) would be correct only if deflating industry prices and values by total costs was the only legitimate way to remove the effect of industry size on estimates of (1). The problem was further discussed by Kliman (2008) and Diaz and Osuna (2008), who, however, reasserted their respective positions[2]. On the other hand, supporters of "Ricardo's 93% Theory of Value" argued that observing physical quantities is not relevant once using input-output data. In that case it is only necessary to assume that physical units of measurement do not change over the period of observation (Tsoulfidis and Paitaridis, 2009, p. 213).

This debate was mainly econometric in nature. However, Kliman (2004, pp. 228-231) offered test results, grounded on Marxian economic theory and again based on national accounts data, that disproved "Ricardo's 93% Theory of Value" also in a cross-sectional setting, once the widespread assumption of a uniform rate of surplus value was embraced. This test was not challenged by the impossibility of observing the physical quantities of output, because no attempt was made to compute unit prices and values as in (2) and (3). The test was instead based on industry gross output, total costs, and the ratio between variable capital and total costs.

The aim of this paper is to use panel data econometrics to shed further light on the issue. First we manipulate equation (2) so as to put in evidence $\frac{p_i}{p_j}$ at time 0, instead of $\frac{q_i}{q_j}$ as in (3). In this setting, we test the above predictions as follows. We start with the hypothesis that $\beta = 1$ and use panel

[2] Diaz and Osuna (2009) extended their critique to price-value deviations.

data econometrics to account for the unobserved heterogeneity induced by the impossibility of controlling for sectoral physical quantities of output and, therefore, for relative unit prices. If the data do not reject this prediction, we will then consider the hypothesis that $\alpha = 0$ on imposing the restriction $\beta = 1$. It is sufficient that only one of these assumptions does not hold to reject the proposition that relative industry values are the main determinants of relative industry prices. We apply our testing procedure to sectoral data for 10 OECD countries over different time periods and aggregation levels.

The rest of this paper is structured as follows. The next Section describes our testing procedure and methods in greater detail. Section 3 discusses our data sources and the way in which we define our variables. Section 4 sets out our results, and Section 5 concludes.

2 Testing procedure and methods

Our testing procedure is as follows. Consider equation (2) at time t

$$\ln \frac{p_{it}}{p_{jt}} = \alpha + \beta \ln \frac{d_{it}}{d_{jt}} + u_{it}$$

Add and subtract from the left hand side $(1 - \beta)$ times the relative output evaluated at base year prices to obtain

$$\ln \frac{p_{it} p_{i0} q_{it}}{p_{jt} p_{j0} q_{jt}} = \alpha + \beta \ln \frac{d_{it} p_{i0} q_{it}}{d_{jt} p_{j0} q_{jt}} + (1 - \beta) \ln \frac{p_{i0} q_{it}}{p_{j0} q_{jt}} + u_{it} \tag{4}$$

This equation can be rewritten as

$$\ln \frac{p_{it}q_{it}}{p_{jt}q_{jt}} = \alpha + (\beta - 1) \ln \frac{p_{i0}}{p_{j0}} + \beta \ln \frac{d_{it}q_{it}}{d_{jt}q_{jt}} + (1 - \beta) \ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}} + u_{it}$$

Bring $\ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}}$ to the left hand side

$$\ln \frac{p_{it}q_{it}}{p_{jt}q_{jt}} - \ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}} = \alpha + (\beta - 1) \ln \frac{p_{i0}}{p_{j0}} + \beta \left(\ln \frac{d_{it}q_{it}}{d_{jt}q_{jt}} - \ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}} \right) + u_{it} \quad (5)$$

Given that the jth good is the numeraire, we drop the j index and define

$$\begin{aligned} y_{it} &\equiv \ln \frac{p_{it}q_{it}}{p_{jt}q_{jt}} - \ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}} \\ x_{it} &\equiv \ln \frac{d_{it}q_{it}}{d_{jt}q_{jt}} - \ln \frac{p_{i0}q_{it}}{p_{j0}q_{jt}} \\ \mu_i &\equiv (\beta - 1) \ln \frac{p_{i0}}{p_{j0}} \\ e_{it} &\equiv u_{it} \end{aligned}$$

Hence we can write (5) as

$$y_{it} = \alpha + \beta x_{it} + \mu_i + e_{it} \quad (6)$$

It is now clear that it is possible to obtain an estimate of β and to test hypotheses about it by applying the well-known one-way error component model in either its fixed effects variant or its random effects one. The two variants differ according to the assumptions concerning μ_i. The fixed effects model considers them as fixed parameters, while the random effects model considers them as random realizations from stochastic processes that are independently and identically distributed with a given variance (Baltagi, 2001, pp. 12-21).

We choose between the two models according to the result of the Hausman test, which is based on the difference between their estimated values of β. The null of this test is that the two estimators produce the same results. Its basis is that fixed effects estimates are consistent but not efficient under both the null hypothesis and the alternative one, while the random effects estimator is not consistent under the alternative hypothesis but it is efficient under the null (Baltagi, 2001, pp. 65-66). This test, however, does not suit all possible datasets, because its underlying assumption that the variance-covariance matrix of the difference between the two estimators is positive definite may not hold in practice. For this reason we supplement it with a Mundlak test (Hsiao, 2001, pp. 50), which is not based on this hypothesis. We further test for serial correlation in e_{it}, by resorting to the Lagrange Multiplier (LM) tests proposed by Baltagi (2001, p. 91 and p. 95). Finally, in the presence of evidence of serial correlation in the residuals, we look for the most suitable specification between an AR(1) and a MA(1) process by resorting to the test proposed by Burke, Godfrey and Termayne (1990) - hereafter referred to as BGT - as illustrated in Baltagi (2001, p. 97).

The above procedure is appropriate when dealing with stationary data.

However, when the time dimension of the panel grows large, a problem of spurious regression may arise (Baltagi, 2001, p. 243). To overcome it, we resort to panel unit root and cointegration testing after Im, Pesaran and Shin (2003), Maddala and Wu (1999), Choi (2001), Pedroni (1999, 2004), Pesaran (2007) and Lewandowski (2008).[3]

If the unit root and cointegration hypotheses are not rejected, we will adopt a panel dynamic least square dummy variables (DLSDV) estimator after Mark and Sul (2003) to test for $\beta = 1$, which, once again, considers μ_i as fixed constants. If no cointegration is not rejected or if series have different orders of integration, this will mean that relative prices and values have no connection in the long-run, which is contrary to "Ricardo's 93% theory of value" irrespective of parameter estimates.

For panels with both short and long time dimensions, once we find evidence of $\beta = 1$, we impose this restriction on the data. In this setting we can test for $\alpha = 0$, because if $\beta = 1$, there will be no omitted variable problem in (3). Finally, note that our results are not invariant to the choice of the numeraire sector. However, rejection for a specific numeraire sector is enough to challenge the generality of "Ricardo's 93% theory of value".

[3] Once again an introduction to these tests is provided in Baltagi (2001, chp. 12).

3 The dataset and definitions of variables

Our data source is the STAN OECD database[4], from which we take the following variables: consumption of fixed capital (CFCC), intermediate inputs in current prices (INTI), gross output in current prices (PROD), gross output in prices for the year 2000 (PRDK), value added in current prices (VALU), the number of employees (EMPE), the number of self-employed (SELF), and labour costs (LABR). We consider the following countries in the time periods stated: Austria from 1976 to 2009, Belgium from 1995 to 2008, the Czech Republic from 1995 to 2009, Denmark from 1970 to 2007, Finland from 1984 to 2004, Greece from 2000 to 2009, Italy from 1980 to 2008, Norway from 1970 to 2007, Slovenia from 2000 to 2009 and Sweden from 1994 to 2008. The precise list of sectors and the level of aggregation varies from country to country depending on data availability. We give preference to the most disaggregated data possible.[5]

We compute y_{it} as follows[6]

$$y_{it} = \ln \frac{PROD_{it}}{PROD_{nt}} - \ln \frac{PRDK_{it}}{PRDK_{nt}}$$

where the n index denotes the numeraire sector.

In order to obtain x_{it} we have first to compute the money value of output

[4] http://www.oecd.org/document/62/0,3746,en_2649_34445_40696318_1_1_1_1,00.html

[5] A list of the sectors considered for each country as well as a list of numeraire sectors is available at the following link: http://www.webalice.it/avaona/List%20of%20sectors.xlsx.

[6] We thank Andrew Kliman for help with the variable definitions.

(MV). Following Kliman (2002, 2004) we proceed as follows. We correct LABR by the wage equivalent for the self-employed (which accounts for the average opportunity cost of not being an employee)

$$LABR'_{it} = LABR_{it}\left(1 + \frac{SELF_{it}}{EMPE_{it}}\right)$$

The aggregate surplus value (S) and rate of surplus value (RSV) are respectively

$$S_{\cdot t} = \sum_i (VALU_{it} - LABR'_{it} - CFCC_{it})$$

$$RSV_{\cdot t} = \frac{S_{\cdot t}}{\sum_i LABR'_{it}}$$

We impose the restriction that sectoral rates of surplus value are all equal to the aggregate one. Therefore sectoral surplus values are

$$S_{it} = RSV_{\cdot t} \cdot LABR'_{it}$$

As a consequence sectoral MVs are

$$MV_{it} = S_{it} + LABR'_{it} + INTI_{it} + CFCC_{it}$$

Note that by construction

$$\sum_i MV_{it} = \sum_i PROD_{it}$$

Finally

$$x_{it} = \ln \frac{MV_{it}}{MV_{nt}} - \ln \frac{PRDK_{it}}{PRDK_{nt}}$$

It is worth recalling that there exist two controversial issues concerning the data to study price-value correlations/deviations. In the first place, the available studies tend to differ regarding whether some sectors should be excluded from the analysis.

Shaikh (1984) omitted the real estate and rental sectors. Ochoa (1989), Steedman and Tomkins (1998), and Tsoulfidis and Maniatis (2002) did not mention any omission. Cockshott et al. (1995, 1997) run different tests once including and excluding the oil sector, whose price could have a sizeable rent portion. Kliman (2002, 2004) was vague regarding the precise industries that were excluded. Zacharias (2006) treated as unproductive expenses those related to financial activities, real estate services, public administration, defence and social security. Instead education and health services were included notwithstanding that they may contain a large share of non-marketed output. Tsoulfidis and Rieu (2006) quoted Shaikh (1998) and they argued that the distinction between productive and unproductive activities is not relevant to the issue under analysis. Tsoulfidis and Mariolis (2007) omitted finance and real estate activities because the concept of output is problematic there. They also omitted public administration and education because the concepts of labour values and prices of production have no meaning in those sectors. Tsoulfidis (2008) did not omit any sector.

Following Dìaz and Osuna (2005-6), among others, we restrict our attention to private activities alone. This is because we impose a uniform rate of surplus value and surplus value requires commodity production within hierarchical organizations to be extracted. However public activities, such as education and health services, for instance, do not generally produce commodities.

The second issue is whether to use either input-output data or national accounts one, which involve different computation methods. There are few studies comparing results across these methods. Tsoulfidis and Paitaridis (2009) relied on input-output data from Canada for the year 1997. They found that empirical results about the deviations of either values or production prices from market prices do not change substantially across different computation approaches. Their critique of Diaz and Osuna (2005-2006) is not theoretical but empirical. They showed that for Canada and for the year 1997 labour values are compatible with the principle of equal profitability in a vertically integrated approach and not once following the one by Diaz and Osuna (2005-2006). However, this was showed for only one country and one year. Vaona (2012) studied the persistence of the deviation of market prices from either production prices or values for different countries and various time periods. His findings are that on computing values and production prices following Kliman (2002, 2004) and Diaz and Osuna (2005-2006, 2007), their deviations from market prices contain a unit root. When using the data by Tsoulfidis (2008) for Japan, instead, there is no unit root, but persistence is nonetheless high. This last result may be due to the short time dimension of the dataset, and from an economic point of view it would anyway point to the fact that the connection between different price sets is loose.

Empirical studies apart, Kliman (2002, p. 301) contended that his approach, though using national accounts data, is theoretically consistent with the one by Ochoa (1984), which both used input-output matrices and is very influential in the literature of reference. Furthermore, suppose to follow Farjoun and Machover (1983) arguing that the equalization of profit rates does not take place, the rate of surplus value has a nearly degenerate distribution and relative prices are determined by relative values. Then, one could not accept the test by Tsoulfidis and Paitaridis (2009) and little would be left to object against Kliman's approach.

More research across different countries and years would be necessary to convincingly dismiss either one or more computation procedures on empirical grounds. We think that the issue of the best computation approach is still open and results are far from being definitive. This would be a research topic in itself.

To have some confidence that our results are not driven by our data or computation method, we estimated (1) . For all the countries under analysis and for all the time periods we can replicate the result that α and β are not statistically different from zero and one respectively (Figures 1 and 2). In all the cases the R^2 of the regressions were close to 0.95. For all these reasons, we believe it legitimate to adopt the computation approach by Kliman (2002, 2004).

4 Results

As mentioned above we use different methods depending on the length of the time span (T) of the data available for each country. Given that spurious regression can generically arise as T grows large (Baltagi, 2001, p. 243 and Entorf, 1997), it is difficult to choose an empirical criterion to single out the countries for which to resort to unit root and cointegration methods. However, Mark and Sul (2003) present an empirical application of their estimation method for a dataset with T=40. For this reason, we consider as long panels those that have a T dimension closer to 40, namely those of Austria, Denmark, Norway and Italy. Short panels are the remaining ones.[7] The results would not change substantially on altering this classification.

4.1 Short panels

Table 1 sets out our results concerning short panels. Both the Hausman and the Mundlak tests always prefer the fixed effects model to the random effects one, with the exception of Greece, where the contrary is the case. LM tests find evidence of serial correlation, and the BGT test points to the AR(1) model rather than to the MA(1) one for the stochastic error. As a consequence, we estimate an AR(1) fixed effects model for all the countries,

[7] Note that the cross-sectional dimension of each panel depends on the number of sectors considered, which varies from country to country and it is reported in the tables below.

but Greece, for which an AR(1) random effects model is implemented. For all the countries except Sweden, the 95% confidence interval does not include the value of 1. Hence we consider the hypothesis that relative values are the main determinants of relative prices as rejected at the 5% level.

Regarding Sweden, we proceed as anticipated above. We impose the restriction that $\beta = 1$ and we check the 5% confidence interval of α' of the following regression

$$y_{it} - x_{it} = \alpha' + \xi_{it}$$

where ξ_{it} is a stochastic error. The estimated value of α' is 0.16 with a p-value of 0.00. Considering each year separately would produce very similar results. For Sweden too, therefore, there is no statistical support for the hypothesis that relative values are the main determinants of relative prices.

4.2 Long panels

Tables 2 to 5 set out our results concerning long panels. For all the countries, except Denmark there is no evidence in support of the model. For Austria and Norway, the series have the same order of integration, but the null of no cointegration is not rejected by the vast majority of the tests. For Italy, the series do not have the same order of integration.

Danish data return controversial results. Once using first generation panel unit root tests, series have the same order of integration, but not when using the test by Pesaran (2007). Suppose, nonetheless, to read this evidence as favorable to "Ricardo's 93% theory of value" and to proceed estimating β. When the panel DLSDV estimator is applied to Danish data - including, in accordance with this method, the fifth, third, second and first leads of Δx_{it} as well as its second and first lags[8] - the estimated value of β is 0.97 with a 95% confidence interval of (0.951,1.005). Hence we find statistical evidence supporting the hypothesis $\beta = 1$. For this reason, we proceed as with Sweden in the previous section. Here, we obtain an estimated value of α equal to 0.03 with a p-value of 0.00. Considering separate regressions for each year, in some instances one finds a value of α statistically not different from 0 at the 5% level (such as for 1970 and 1971), but for some other years its value is highly statistically significant. Also for Denmark, then, there is no statistical support for the hypothesis that relative values are the main determinants of relative prices.

5 Conclusions

This paper has furnished new empirical insights into price-value correlations. There has recently been a debate on whether it is possible to offer tests for the proposition that relative values are the main determinants of relative prices. We have shown that panel data econometrics can offer a test for this proposition, overcoming the problems of indeterminacy that may arise

[8] Leads and lags are selected according to their significance, by dropping regressors insignificant at the 5% level. All the remaining regressors are significant at the 1% level, except for the constant, which has a p-value of 0.06.

in cross-sectional estimates. The results obtained for 10 OECD countries do not provide support for "Ricardo's 93% Theory of Value".

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Table 1 - Fixed and Random effects estimates and model specification tests for various OECD countries

Country	Timespan	N. sectors	β	95% confidence interval		Hausman test (p-values)	Mundlak test (p-values)	LM test (p-values)	BGT Test (p-values)
Fixed effects model									
Belgium	14	46	0.51	0.45	0.57	0.00	0.00	0.00	0.67
Czech	15	49	0.67	0.64	0.70	0.00	0.00	0.00	0.87
Finland	21	47	0.75	0.70	0.79	0.00	0.00	0.00	0.79
Slovenia	10	45	0.22	0.18	0.27	0.00	0.00	0.00	0.67
Sweden	15	31	0.94	0.86	1.02	-	0.00	0.00	0.76
Random effects model									
Greece	10	49	0.00	-0.04	0.05	0.11	0.10	0.00	0.25

Notes. The Hausman test is distributed as a χ squared with 1 degree of freedom. Its null is that the random effects model is preferable to the fixed effects one. The same null applies in the Mundlak test, which has an F distribution with degrees of freedom equal to the number of regressors and the number of observations minus twice the number of regressors plus one. The LM test is distributed as a N(0,1) for the number of time periods going to infinity. See Baltagi (2001), pp. 94-95. Its null is the absence of serial correlation. For Greece only, the LM test is instead the one presented by Baltagi (2001), pp. 90-91, and it is distributed as a χ squared with 2 degrees of freedom for the number of cross-sectional units going to infinity. Its null is the absence of serial correlation and that the variance component due to sectoral specificieties is equal to zero. The BGT test is the Burke, Godfrey and Termayne (1990) test illustrated by Baltagi (2001), pp. 98-99. It is distributed as a N(0, 1) for a large number of sectors. Its null is that the error process of the estimated model can be modelled as an AR(1) rather than an MA(1).

Table 2 - Panel unit root and cointegration tests, 45 Austrian sectors from 1976 to 2009

Panel A: Panel Unit Root Tests. Null hypothesis: all the series have a unit root

	Y	∆Y	X	∆X
Im, Pesaran and Shin W-stat	2.91 (0 to 3)	-24.11a (0 to 7)	0.54 (0 to 5)	-32.0071a (0 to 1)
ADF - Fisher Chi-square	50.61 (0 to 3)	668.60a (0 to 7)	81.68 (0 to 5)	900.973a (0 to 1)
PP - Fisher Chi-square	50.00 (0 to 3)	683.81a (0 to 7)	73.55 (0 to 5)	931.477a (0 to 1)
Pesaran - CADF	-1.32 (1)	-3.24a (1)	-1.57 (1)	-4.00a (1)

Panel B: Panel Cointegration Tests. Null hypothesis: no cointegration

Within dimension test statistics		Between dimension test statistics	
Panel v-statistic	-2.626027b	Group rho-statistic	1.695307
Panel rho-statistic	1.139610	Group PP-statistic	0.657860
Panel PP-statistic	0.154986	Group ADF-statistic	0.427007
Panel ADF-statistic	0.275928		

Notes: variables expressed in natural logarithms. Panel unit root tests include intercepts. Automatic lag length selection was performed on the basis of the Schwarz information criterion. Of the seven cointegration tests, the panel v-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration, whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration. All the cointegration tests were carried out without including a trend. For lag selection in the cointegration tests we used the Schwarz information criterion. 1 percent significance level denoted by ‘‘a’’, 5 per cent significance denoted by "b" and 10 per cent by "c".

Table 3 - Panel unit root and cointegration tests, 35 Danish sectors from 1970 to 2007

Panel A: Panel Unit Root Tests. Null hypothesis: all the series have a unit root

	Y	ΔY	X	ΔX
Im, Pesaran and Shin W-stat	8.32 (0 to 2)	-32.67a (0 to 1)	2.49 (0 to 4)	-35.09a (0 to 3)
ADF - Fisher Chi-square	15.69 (0 to 2)	866.83a (0 to 1)	46.62 (0 to 4)	928.40 (0 to 3)
PP - Fisher Chi-square	13.08 (0 to 2)	903.82a (0 to 1)	43.75 (0 to 4)	1002.41 (0 to 3)
Pesaran - CADF	-1.56 (4)	-2.53a (3)	-2.00a (1)	-3.09a (3)

Panel B: Panel Cointegration Tests. Null hypothesis: no cointegration

Within dimension test statistics		Between dimension test statistics	
Panel v-statistic	2.687805b	Group rho-statistic	-4.86a
Panel rho-statistic	-6.942717a	Group PP-statistic	-5.46a
Panel PP-statistic	-5.782381a	Group ADF-statistic	-5.87a
Panel ADF-statistic	-6.291381a		

Notes: variables expressed in natural logarithms. Panel unit root tests include intercepts. Automatic lag length selection was performed on the basis of the Schwarz information criterion. Of the seven cointegration tests, the panel v-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration, whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration. All the cointegration tests were carried out without including a trend. For lag selection in the cointegration tests we used the Schwarz information criterion. 1 percent significance level denoted by ‘‘a’’, 5 per cent significance denoted by "b" and 10 per cent by "c".

Table 4 - Panel unit root tests, 24 Italian sectors from 1970 to 2007

Null hypothesis: all the series have a unit root

	Y	ΔY	X	ΔX
Im, Pesaran and Shin W-stat	-1.70b (0 to 6)	-14.51a (0 to 6)	1.21 (0)	-20.5813a (0)
ADF - Fisher Chi-square	88.64a (0 to 6)	294.72a (0 to 6)	37.21 (0)	407.743a (0)
PP - Fisher Chi-square	163.88a (0 to 6)	327.27a (0 to 6)	50.01 (0)	417.104a (0)
Pesaran - CADF	-2.23a (1)	-3.34a (1)	-1.60 (1)	-3.28a (1)

Notes: variables expressed in natural logarithms. Panel unit root test includes intercept. 1 percent significance level denoted by ‘‘a’’, 5 per cent significance denoted by "b" and 10 per cent by "c".

Table 5 - Panel unit root and cointegration tests, 42 Norwegian sectors from 1970 to 2007

Panel A: Panel Unit Root Tests. Null hypothesis: all the series have a unit root

	Y	ΔY	X	ΔX
Im, Pesaran and Shin W-stat	6.46 (0 to 6)	-19.56a (0 to 5)	4.37 (0 to 9)	-35.63a (0 to 5)
ADF - Fisher Chi-square	39.68 (0 to 6)	529.81a (0 to 5)	53.39 (0 to 9)	1024.83a (0 to 5)
PP - Fisher Chi-square	32.12 (0 to 6)	599.37a (0 to 5)	70.29 (0 to 9)	1089.11a (0 to 5)
Pesaran - CADF	-0.76 (1)	-3.61a (1)	-1.18 (2)	-4.41a (1)

Panel B: Panel Cointegration Tests. Null hypothesis: no cointegration

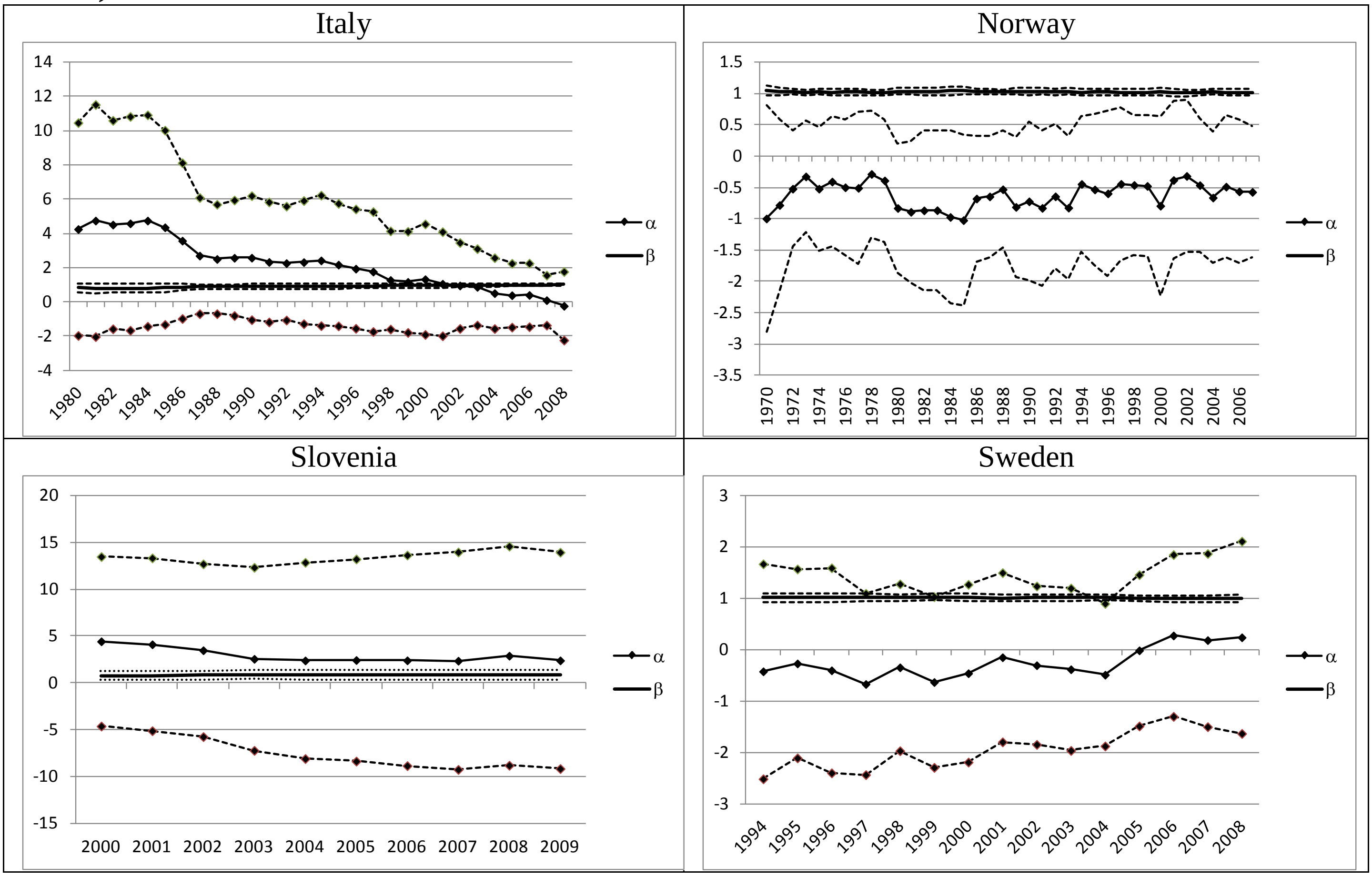
Within dimension test statistics		Between dimension test statistics	
Panel v-statistic	-0.161473	Group rho-statistic	1.288781
Panel rho-statistic	-0.718028	Group PP-statistic	-0.971838
Panel PP-statistic	-1.828059	Group ADF-statistic	-1.189264
Panel ADF-statistic	-1.873610		

Notes: variables expressed in natural logarithms. Panel unit root tests include intercepts. Automatic lag length selection was performed on the basis of the Schwarz information criterion. Of the seven cointegration tests, the panel v-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration, whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration. All the cointegration tests were carried out without including a trend. For lag selection in the cointegration tests we used the Schwarz information criterion. 1 percent significance level denoted by ‘‘a’’, 5 per cent significance denoted by "b" and 10 per cent by "c".

Figure 1 - Coefficient estimates of cross-section regressions of aggregate prices over aggregate values, selected countries

Note: dotted lines represent 95% confidence intervals.

Figure 2 - Coefficient estimates of cross-section regressions of aggregate prices over aggregate values, selected countries



Note: dotted lines represent 95% confidence intervals.