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Invited article

The price and income elasticities of the top clothing exporters: Evidence from a panel data analysis

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A B S T R A C T

This paper studies the main export function features of twelve top clothing exporters (China, Hong Kong, France, Germany, India, Indonesia, Italy, Netherlands, Spain, Turkey, UK and USA) in the period between 1992 and 2011. Price and income elasticities are estimated for each economy using a panel data approach, after controlling for nonstationarity, cointegration and Granger causality. Rolling regressions are also performed, and show the existence of some elasticities instability over time, fundamentally related to the profound economic and institutional changes affecting the clothing trade in the period under consideration. The analysis suggests that most advanced economies, including Hong Kong, changed their position in the global value chain towards an "organizational" role. China confirms its leadership in clothing exports although its rising price elasticity sounds a warning with regard to future prospects.

1. Introduction

The analysis of the size and time stability of an economy's export elasticities is of central importance in empirical trade studies (see, among others, Arize, 2001; Aziz & Li, 2008; Hooper, Johnson, & Marquez, 2000; Sharma, 2001; Thorbecke, 2010), given their fundamental role in terms of growth performance, international competitiveness, balance of payments equilibrium and industrial policy decisions. In fact, as noted by Aziz and Li (2008), if export price and income elasticities are low, changes in external conditions or in the exchange rate are unlikely to have much impact on an economy's growth or its current account dynamics. However, if they are not stable, little can be said about how an economic system might respond to such changes. In empirical trade literature, several papers focusing on Asian economies study the size and stability of export function parameters for the whole system (Arize, 1990, 2001; Aziz & Li, 2008; Lucas, 1988; Muscatelli, Srinivasan, & Vines, 1992). The most widely studied case is China, given its increasing importance in international trade especially after its entry into the World Trade Organization. In particular, Yao, Tian, and Su (2013) estimate China's export elasticities for the time period 1992–2006 and find that its outstanding performance is due to the joint influence of a very high income elasticity (2.34) and a surprisingly low price elasticity (−0.65). However, Aziz and Li (2008) find that in the same period price and

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Table 1
Top 15 Clothing exporters: export values (Billion of USD) and market shares in 2011.

	Export value	Market share (%)
China	153.8	37.28
Hong Kong	24.5	5.94
Italy	23.2	5.64
Bangladesh	19.9	4.83
Germany	19.6	4.76
India	14.4	3.48
Turkey	13.9	3.38
Vietnam	13.2	3.19
France	11.0	2.67
Spain	9.2	2.24
Belgium	9.0	2.20
Netherlands	8.2	2.00
Indonesia	8.0	1.95
United Kingdom	6.6	1.59
United States	5.2	1.27
Total	340.0	82.42

Source: Authors' elaboration on WTO data.

income elasticities were not stable, but increased over time, because of changes in the composition and degree of sophistication of exports.

All the above mentioned studies use aggregate data and real effective exchange rates (REERs) to estimate the basic parameters of the export functions. Some studies highlight the importance of using disaggregated sectoral data in order to obtain specific trade elasticities at the industry level, but still adopt aggregate REER indexes in their estimations. Finally, others make use of industry-specific REERs (Dai & Xu, 2013): this approach is the most appropriate from both a theoretical and an econometric standpoint, as confirmed by the fact that a significant relationship between sectoral exchange rates and exports can be identified here (Dai & Xu, 2013), unlike many studies using aggregate REERs.

Sectoral elasticities have the further advantage of allowing comparisons between different economies in specific industrial sectors, in order to shed light on their positioning in the global market and avoid problems connected to the composition fallacy (for details, see Mayer, 2003). Price elasticities can thus indicate the relative strength of an economy's production, since it is likely that high-quality goods will exhibit an inelastic foreign demand and vice versa. On the other hand, income elasticities, closely correlated with the export growth rate, reflect the non-price competitiveness of an economy, and are influenced by factors such as export composition by goods and destination markets, embodied technology, marketing strategies and promotion, distribution services, financial assistance to exporters and so on.

The aim of this paper is to estimate the export price and income elasticities for the world's top exporters in the clothing industry. This sector was chosen as it has been the driving force behind the impressive export performance of many Asian economies, including China, Hong Kong, Bangladesh, India, Vietnam and Indonesia. This is particularly true for China, the leading exporter in 2011 with a 37.3% share of world exports (see Table 1). The other five economies are among the top ten exporters and together account for another 19.4% of world clothing exports. Clothing is also a low-tech industry which still plays an important role in value added creation and trade in many advanced economies[2], despite recent strong competition from lower-wage nations. Finally, clothing also carries the quality reputation of an economy and its productive system.

The contribution of this paper is threefold. First, price and income elasticities for the top exporters in the last twenty years are estimated and compared. Second, these elasticities are re-estimated for different sub-periods in order to verify whether they are stable over time. To the best of our knowledge, this is the first study providing such estimates for twelve top world clothing exporters over a time-span of twenty years. Third, our econometric analysis allows insights into strengths, weaknesses and prospects for the clothing industry derived by jointly considering market shares, export average unit values (AUVs) and price and income elasticities for each sample economy. In particular the paper provides the theoretical and empirical background to assess whether investing in a traditional industry like clothing is still profitable and thus sustainable, especially in the case of advanced economies. This conclusion would appear to be at variance with recent literature on industrial policy, which mainly suggests discouraging investment in traditional sectors and promoting R&D expenditure in high-tech industries (see, for example, OECD, 2010; European Commission, 2010), because of their higher productivity growth rates (Fagerberg, 2000; Grossman & Helpman, 1991).

Adopting an original methodology, we use an appropriate measure of relative export prices, obtained as the ratio between each economy's disaggregated export unit values and the AUVs of the whole sample. We also examine issues that have been little investigated in literature in the field, such as the order of integration of trade variables, their possible cointegration in the long run and the direction of Granger causality between them. Our empirical analysis uses a panel data approach, which

[2] See Table A1 in Appendix A for the share of clothing in total manufacturing exports for the twelve economies considered in the analysis.

allows us to use disaggregated AUVs and obtain better results in the study of causality with respect to time series, as demonstrated in many empirical papers (see, among others, Hsiao & Hsiao, 2006).

Rolling-period regressions, which are widely used in empirical papers focused on international trade, especially for economies characterized by profound structural changes (see for example Aziz & Li, 2008), are also performed in order to check the stability of the relevant long-run relationships. If rolling-regression parameters prove to be in line with the long-run estimates, we can obviously conclude that long-run relationships are stable in time. Otherwise, long-run results, even when statistically significant on the basis of standard econometric tests, should be treated with caution in the presence of profound economic and institutional changes. Our findings indeed show that income elasticities are slightly decreasing over time for most economies, while price elasticities are generally stable, with the exception of the largest Asian exporters (China, India, Hong Kong and Turkey) in the most recent time-spans. This result has important implications, especially in terms of economic policy. Long-run elasticities are obviously statistical averages of shorter-period rolling elasticities, but policymakers should perhaps use long-run estimates with caution and pay instead more attention to the latest rolling window parameters when evaluating the response of their economies to external shocks or designing industrial strategies.

The rest of the paper is structured as follows. Section 2 describes the data, in the context of the economic and institutional changes in clothing exports in the last two decades, and introduces the model specification to be used in the subsequent empirical analysis. Section 3 presents the testing framework, which includes unit roots, cointegration tests and panel Granger-causality analysis. Section 4 reports the estimated long-run export price and income elasticities of the considered economies, and proposes a stability analysis of these parameters. Section 5 analyses the estimation results, together with AUV and market share dynamics, to discuss the export performance and prospects of the sample economies, as well as policy implications. Section 6 concludes.

2. Data and model specification

2.1. An overview of clothing export dynamics in a changing institutional environment

Our analysis initially considered the disaggregated clothing export data for the top fifteen economies (Bangladesh, Belgium, China, Hong Kong, France, Germany, India, Indonesia, Italy, Netherlands, Spain, Turkey, UK, USA and Vietnam) in the period between 1992 and 2011. These economies together account for 82.4% of the total value of world clothing exports in 2011 (see Table 1).

The data, retrieved from the UN Comtrade database, are disaggregated at the 4 digit level of the Standard International Trade Classification (Rev. 3), which implies the consideration of 37 distinct goods[3].

Three countries (Bangladesh, Belgium and Vietnam) were excluded from the sample because of incomplete records in the period considered. Our final analysis is therefore carried out on the twelve remaining economies (China, Hong Kong, France, Germany, India, Indonesia, Italy, Netherlands, Spain, Turkey, UK, USA) accounting for 72.2% of world clothing exports in 2011. The available data are organized to form 12 panel datasets, one for each economy. Each dataset thus consists of 740 observations relative to 37 goods at the 4 digit level in the 20-year time period considered.

Since this period has been characterized by far-reaching economic and institutional changes, it is interesting to analyse the export volume dynamics of the sample exporters in more detail, as in Fig. 1.

China, Hong Kong and the USA stand out as the most glaring examples of economic systems experiencing a structural change in their export pattern. In particular, since the beginning of the new millennium China has shown an exponential increase in foreign sales, which in spite of the global crisis of 2008–2009 brought its market share up from slightly below 30% to over 52%. This is obviously linked to the fact that in 1995 the General Agreement on Tariffs and Trade (GATT) Uruguay Round came into effect, and brought the textile and clothing sectors under the jurisdiction of the World Trade Organisation (WTO), which China joined in 2001. The Agreement on Textile and Clothing (ATC) allowed a gradual dismantling of the quotas that existed under the 1974 Multi Fibre Arrangement (MFA), a process which ended in 2005. At the beginning of that year, exports from China to the West grew rapidly leading the USA and European countries to restrict the rate of growth of Chinese imports to 7.5% per year until 2008.

This change in China's trade performance impacted on all major clothing exporters and especially Hong Kong and the USA. At the end of the 1990s, Hong Kong lost its importance in production and became instead a regional trading and sourcing hub, bringing together buyers, sellers and input suppliers (Van Grunsven & Smakman, 2001). In fact, the emergence of China as a leading exporter led to a new specialization model (Bair & Gereffi, 2003), whereby leading retailers and branders in more advanced economies (such as the USA) retain direct control over design, marketing and retailing, and fix product standards for manufacturing suppliers operating in less developed economies (Gereffi, 1994, 1999).

The late 1990s were also a turning point for Turkey and Indonesia. Turkey, in particular, after shifting from an import substitution to an export-led growth strategy in the 1980s, strengthened its association with the European Union and obtained preferential supplier status. In 1996 the establishment of a Custom Union Agreement removed all restrictions on access to the European market (Neidik & Gereffi, 2006). In Indonesia, export-oriented policies implemented in the mid-1980s allowed clothing companies to take advantage of changes occurring in the industrial sector (Dicken & Hassler, 2000). On the

[3] The list of goods is reported in Table A2 in Appendix A.

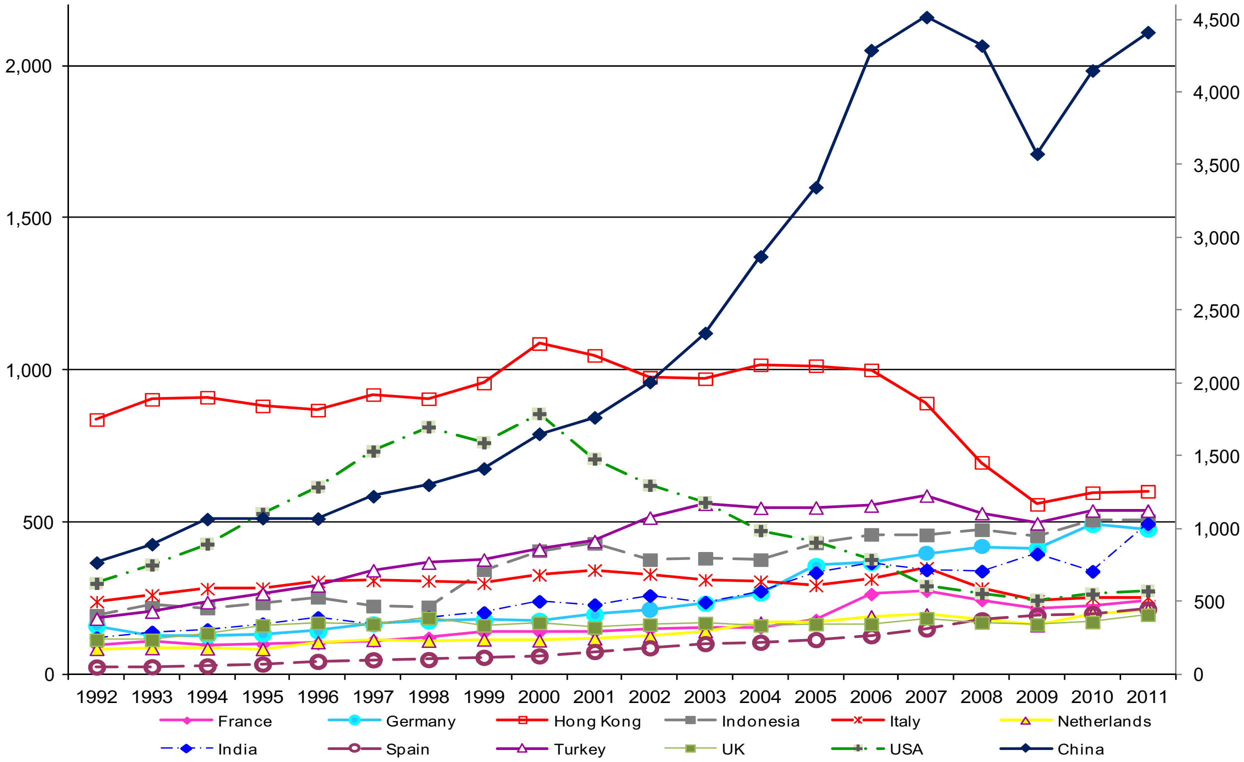

Fig. 1. Export volumes 1992–2011. *Source:* Authors' calculations on UN Comtrade data. *Notes:* All exporters' volumes are measured on the left vertical axis with the exception of China's volumes, measured on the right axis.

other hand, Fig. 1 confirms the claim (Alessandrini, Fattouh, Ferrarini, & Scaramozzino, 2011; Gupta, Hasan, & Kumar, 2008; Kochhar, Kumar, Rajan, Subramanian, & Tokatlidis, 2006; Sen, 2009) that in India the overall impact of tariff reduction appears to have favoured skill-intensive and large scale industries, rather than labour-intensive manufacturing. This sector has not been able to benefit from the reforms or exploit the growing integration with the rest of the world.

Finally, European export markets were challenged very quickly by the ATC, despite its prolonged ten-year adjustment process (Pickles & Smith, 2011). In fact, in the overall period 1992–2011 export volumes generally increased, but market shares mainly fell, with the exception of Spain and Germany, showing, respectively, a substantial and a slight improvement, as shown in detail in Table 2. More specifically, Table 2 reports for each sample economy the descriptive statistics for the variables of interest: export volumes, market shares and average unit values (AUVs), at the beginning and at the end of the time span, together with their average value over the whole period[4,5]. The most interesting case, however, is Italy, where market share in volume more than halves over the period, while the share in value (USD) remains relatively high. This is thanks to the significant increase in AUVs, which practically double, and rise much faster than in the other economies.

2.2. *Model specification*

For every economy we estimate a standard panel data export function, specified as follows:

$$\ln X_{it} = \alpha_i + \beta_i \ \ln \ \mathrm{RP}_{it} + \gamma_i \ \ln \ \mathrm{GDPW}_t + \varepsilon_{it} \tag{1}$$

where i and t refer to the ith good and the tth year, respectively, with $i = 1,\ldots,N$ and $t = 1,\ldots,T$. X_{it} is the yearly export volume for each of the 37 goods; RP_{it} is the yearly relative export price of each good; GDPW_t is the annual world GDP in constant 2005 USD, which is invariant for each cross-section[6]. An innovative feature of our estimations is that the relative price is

[4] Hereafter, export market shares indicate the ratios between the export volume of each economy and the total export volume of the 12 economies considered in the analysis.

[5] The export unit values for each good and economy are computed by dividing exports values by their volumes. Similarly, average export unit values for each good of the whole sample economies are obtained by dividing total export values of each good by their total volumes.

[6] We differ from previous literature in that we do not study the export elasticities of any economy with respect to specific destination markets but rather the elasticities of top clothing exporters with respect to the whole world. This is because we are interested in comparing not the differences in the export functions with respect to destination markets, following a substantial within-economy approach, but rather the differences with regard to origin exporting economies, in an across-economies approach.

Table 2
Descriptive statistics: export volumes, market shares and average unit values, beginning-of-period, end-of-period and average-of-period.

	Export volumes (tons)			Market shares (% of export volumes)			Average unit values (USD)		
	1992	2011	Average	1992	2011	Average	1992	2011	Average
China	766,578	4,411,045	2,401,827	24.67	52.43	37.39	21.79	34.86	24.10
France	94,758	241,942	161,746	3.05	2.88	2.70	55.58	45.64	46.05
Germany	158,99	476,753	260,227	5.12	5.67	4.22	52.54	42.13	43.79
Hong Kong	836,787	600,745	881,465	26.93	7.14	16.60	23.97	40.79	27.96
India	119,73	494,487	257,506	3.85	5.88	4,27	25.97	29.63	26.48
Indonesia	193,199	504,879	358,149	6.22	6.00	6.07	16.38	15.93	12.99
Italy	238,211	251,866	293,345	7.67	2.99	5.42	51.40	92.31	58.15
Netherlands	83,144	212,011	137,225	2.68	2.52	2.93	32.51	41.20	28.43
Spain	23,194	213,990	94,751	0.75	2.54	1.45	30.53	42.85	35.91
Turkey	182,137	537,145	425,891	5.86	6.38	7.17	22.95	25.97	20.73
UK	112,924	194,793	160,879	3.63	2.32	2.91	32.40	34.04	29.42
USA	298,117	273,195	494,719	9.59	3.25	9.49	14.13	19.12	13.00

Source: Authors' elaboration on WTO data.

obtained as the ratio between the export unit value of each selected economy for every good i at time t and the average export unit value of all economies considered for the same good and time. We believe that this measure corresponds to the actual price much more precisely than the usually employed aggregate measures such as CPI-based REERs etc. All variables are transformed into natural logarithms and labelled $\ln X_{it}$, $\ln \mathrm{RP}_{it}$ and $\ln \mathrm{GDPW}_t$. The coefficients β_i and γ_i are the clothing export price and income elasticities, respectively. The former are expected to be negative, and the latter positive. The α_i are the intercepts for each good and the ε_{it} the error terms.

3. A causality analysis between export volumes, relative prices and world income

3.1. Panel unit root tests

As a first step, it is necessary to check whether each variable of interest is stationary. For this purpose, it is common practice in the literature to perform several panel unit root tests, given the shortcomings of any single test with regard to sample size and power properties. Hence we propose three different panel unit root tests: the Breitung (2000) and the Hadri (2000) tests, which assume homogeneity among each cross section, and a more recent test developed by Pesaran (2007). The null hypothesis of these unit root tests is that all series contain a unit root, with the exception of the Hadri test, whose null hypothesis is that all panels are stationary. The Hadri unit root test is performed to confirm or reject any conclusion based on the null hypothesis of nonstationarity. Moreover, the Pesaran unit root test is more powerful because it takes into account the possible presence of cross-sectional dependence in heterogeneous panels. As underlined by Pesaran (2007), this test is appropriate even in the case of very small sample sizes (i.e., when N and T are equal to 10). Finally, two popular time series unit root tests are employed to assess the order of integration of GDPW_t: the augmented Dickey–Fuller (ADF) and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) tests (Said & Dickey, 1984; Kwiatkowski, Phillips, Schmidt, & Shin, 1992, respectively). This choice is explained by the nature of this variable, which is a time series, invariant for each cross-section.

Tables 3a–3c show the results of the above specified unit root tests. Export volumes ($\ln X_{it}$) are found to be $I(1)$ for all economies, with the UK as the only exception in the case of the Breitung test, which rejects the null hypothesis of nostationarity. As regards relative prices ($\ln \mathrm{RP}_{it}$) the Breitung test suggests stationarity for all economies with the exception of China and Indonesia[7]. However, both the Hadri and Pesaran tests show evidence of nonstationarity for all twelve panels of exporters. For these reasons, also $\ln \mathrm{RP}_{it}$ should be properly considered as $I(1)$. Finally, the time series unit root tests confirm the nonstationarity of world GDP ($\ln \mathrm{GDPW}_t$) as expected.

3.2. Panel cointegration tests

In order to investigate the existence of a long-run relationship between the considered variables, the Pedroni (1999) and Kao (1999) cointegration tests are computed for the 12 panel datasets. Both tests are based on the Engle–Granger (1987) two-stage cointegration test framework. The Pedroni test allows for heterogeneity across cross-sections in terms of intercept and trend coefficients. After estimating Eq. (1), we test the stationarity of residuals, which will be $I(1)$ under the null hypothesis of no cointegration in a heterogeneous panel. For this purpose, Pedroni (1999, 2000) suggests two types of residual-based tests: for the first type, four tests (panel-$\boldsymbol{v}$, panel-ρ, panel-pp and panel ADF-statistics) are based on pooling the residuals of the regression along the within-dimension of the panel (panel tests); for the second type, three tests (group-ρ, group-pp, and group ADF-statistics) are based on pooling the residuals of the regression along the between-dimension of

[7] The null hypothesis is not rejected at the 10% significance level in the case of Germany.

Table 3a
Panel unit root test statistics for the variable $\ln X_{it}$.

	Breitung	Hadri	Pesaran without trend	Pesaran with trend
China	6.74 (0.88)	63.89 (0.00)	−0.70 (0.24)	0.40 (0.65)
France	2.21 (0.99)	53.86 (0.00)	0.77 (0.78)	3.32 (1.00)
Germany	4.86 (1.00)	62.48 (0.00)	0.01 (0.50)	0.99 (0.84)
Hong Kong	1.39 (0.91)	49.09 (0.00)	3.69 (1.00)	0.89 (0.81)
India	2.85 (0.99)	42.19 (0.00)	−0.91 (0.18)	−0.92 (0.18)
Indonesia	−0.14 (0.44)	49.50 (0.00)	−0.54 (0.29)	0.34 (0.64)
Italy	0.24 (0.59)	36.43 (0.00)	2.90 (0.99)	3.08 (0.99)
Netherlands	1.02 (0.85)	43.12 (0.00)	0.07 (0.53)	0.41 (0.66)
Spain	7.62 (1.00)	63.24 (0.00)	−0.39 (0.35)	1.54 (0.94)
Turkey	3.62 (0.99)	43.72 (0.00)	0.69 (0.75)	2.71 (0.99)
UK	−2.54 (0.00)	32.47 (0.00)	−0.72 (0.24)	−0.42 (0.34)
USA	−1.08 (0.14)	42.52 (0.00)	1.04 (0.85)	1.98 (0.97)

Source: Authors' calculations on UN Comtrade data. *Notes:* The lambda-statistics, *z*-statistics and the standardized *Zt*-bars are reported for the Breitung (2000), Hadri (2000) and Pesaran (2007) unit root tests, respectively; *p*-values in parentheses; the Breitung and Hadri tests are calculated by including the intercept in the test equations, while the Pesaran test is computed by adding also a time trend; the null hypothesis for all tests is "Panels contain unit roots" with the exception of the Hadri test, whose null hypothesis is "All panels are stationary".

Table 3b
Panel unit root test statistics for the variable $\ln RP_{it}$.

	Breitung	Hadri	Pesaran without trend	Pesaran with trend
China	1.21 (0.89)	31.93 (0.00)	−0.87 (0.19)	−0.58 (0.28)
France	−2.61 (0.00)	36.35 (0.00)	−0.11 (0.46)	−0.58 (0.28)
Germany	−1.52 (0.06)	47.39 (0.00)	−0.24 (0.41)	−0.41 (0.34)
Hong Kong	−1.74 (0.04)	38.05 (0.00)	−0.36 (0.36)	−1.17 (0.12)
India	−3.34 (0.00)	36.02 (0.00)	0.86 (0.80)	1.62 (0.95)
Indonesia	−0.01 (0.49)	44.34 (0.00)	0.47 (0.68)	0.69 (0.76)
Italy	−5.10 (0.00)	23.72 (0.00)	−0.31 (0.38)	0.16 (0.56)
Netherlands	−3.81 (0.00)	38.09 (0.00)	−0.82 (0.20)	−0.84 (0.20)
Spain	−3.20 (0.00)	31.21 (0.00)	1.83 (0.97)	3.43 (1.00)
Turkey	−3.23 (0.00)	20.92 (0.00)	−0.25 (0.40)	0.19 (0.58)
UK	−6.49 (0.00)	21.43 (0.00)	−1.03 (0.15)	−0.34 (0.37)
USA	−3.51 (0.00)	30.58 (0.00)	−1.22 (0.11)	−0.07 (0.47)

Source: Authors' calculations on UN Comtrade data. *Note:* See Table 3a.

Table 3c
Unit root tests for the variable $\ln GDPW_t$.

ADF	KPSS
0.34 (0.97)	0.61 [0.46]

Source: Authors' calculations on UN Comtrade data. *Notes:* The *t*-statistics and LM-statistics are reported for the augmented Dickey–Fuller (ADF) and the Kwiatkowski–Phillips–Schmidt–Shin (KPSS) unit root tests; *p*-values and asymptotic critical values in parentheses and brackets, respectively; an asymptotic critical value of 0.46 corresponds to the 5% significance level; the ADF and the KPSS unit root tests are calculated by including the intercept in the test equations; the null hypothesis is "$\ln GDPW_t$ has a unit root" for the ADF test and "$\ln GDPW_t$ is stationary" for the KPSS test.

the panel (group tests). In both cases, the hypothesized cointegrating relationship is estimated separately for each panel member and the resulting residuals are then pooled in order to conduct the panel tests[8]. In the case where the seven statistics lead to different outcomes, we follow a common practice in literature (e.g. Bottasso, Castagnetti, & Conti, 2013; Lee & Chang, 2008; Narayan, Smyth, & Prasad, 2007) in assuming that the null hypothesis of no cointegration is rejected if at least four statistics yield evidence of cointegration. Moreover, in our test assessment, we consider that the panel-ADF and group-ADF tests have better small-sample properties than other tests, so that they are more reliable (Pedroni, 1999).

The Kao test follows a similar approach but allows for specific intercepts for each cross-section and homogenous coefficients in the first stage, thus implying heterogeneity in intercepts α_i and homogeneity in β_{1t} and γ_{1i} and all coefficient trends to be zero.

[8] For the within dimension, weighted statistics have also been calculated. They have not been reported in the tables since they confirm the results of unweighted statistics. Estimates are available on request.

Table 4 shows that both the Pedroni and Kao tests reject the null hypothesis of no cointegration in eight out of twelve economies. For the remaining four, one test confirms cointegration while the other one does not. In particular, as far as Hong Kong and Turkey are concerned, the Pedroni test cannot reject the null hypothesis of no cointegration but the Kao test does. On the contrary, in the case of India and Germany, the Pedroni test confirms cointegration, while the Kao test does not. On the whole, since for each economy at least one test confirms cointegration, the presence of a long-run relationship between the variables is assumed.

3.3. Panel Granger causality tests

Given the existence of a cointegration relationship, the next step is to determine the direction of causality between the variables. In particular, since we are interested in studying price and income elasticities, we need to find evidence of a long-run causality from prices and income to export volumes. Panel Granger causality is tested following the two-step Engle–Granger causality procedure (Engle & Granger, 1987). First, we apply the panel Mean Group (MG) estimator proposed by Pesaran and Smith (1995) to the previously specified Eq. (1). The MG estimator is designed for "moderate-T, moderate-N" macro panels, where moderate typically means about 15 time series/cross-section observations. It is part of the panel time-series (or nonstationary panel) literature, which applies in the presence of unit roots, cross-section dependence and parameter heterogeneity[9,10].

The second step consists of building a Granger-causality model with a dynamic error correction term (Holtz-Eakin, Newey, & Rosen, 1988). For this purpose, it is necessary to incorporate the lagged residuals of Eq. (1) into the following dynamic error correction model:

$$\Delta \ln X_{it} = \alpha_i^X + \sum_{l=1}^{p} \theta_{il}^X \Delta \ln X_{i,t-l} + \sum_{m=1}^{q} \eta_{im}^X \Delta \ln \mathrm{RP}_{i,t-m} + \sum_{n=1}^{r} \mu_{in}^X \Delta \ln \ \mathrm{GDPW}_{t-n} + \omega_i^X \mathrm{ECT}_{i,t-1} + u_{it} \tag{2a}$$

$$\Delta \ln \ \mathrm{RP}_{it} = \alpha_i^{\mathrm{RP}} + \sum_{l=1}^{p} \theta_{il}^{\mathrm{RP}} \Delta \ln \ \mathrm{RP}_{i,t-l} + \sum_{m=1}^{q} \eta_{im}^{\mathrm{RP}} \Delta \ln X_{i,t-m} + \sum_{n=1}^{r} \mu_{in}^{\mathrm{RP}} \Delta \ln \ \mathrm{GDPW}_{t-n} + \omega_i^{\mathrm{RP}} \mathrm{ECT}_{i,t-1} + v_{it} \tag{2b}$$

$$\Delta \ln \ \mathrm{GDPW}_{t} = \alpha_i^{\mathrm{GDPW}} + \sum_{l=1}^{p} \theta_{il}^{\mathrm{GDPW}} \Delta \ln \ \mathrm{GDPW}_{t-l} + \sum_{m=1}^{q} \eta_{im}^{\mathrm{GDPW}} \Delta \ln X_{i,t-m} + \sum_{n=1}^{r} \mu_{in}^{\mathrm{GDPW}} \Delta \ln \ \mathrm{RP}_{i,t-n} + \omega_i^{\mathrm{GDPW}} ECT_{i,t-1} + \xi_{it} \tag{2c}$$

where $\mathrm{ECT}_{i,t-1}$ is the lagged residual derived from the long-run cointegrating relationship in Eq. (1). The parameters θ, η and μ in Eqs. (2a)–(2c) are the short-run adjustment coefficients. The parameters ω_i, on the other hand, indicate the speed of adjustment of the model to the long-run equilibrium. The sign of ω_i^X is expected to be negative, in order to assure convergence.

The disturbance terms u_{it}, v_{it} and ξ_{it} are uncorrelated and with zero mean. Δ indicates the first differences of the variables and p, q and r the lag length, usually determined by the Akaike or the Schwarz Information Criteria. In this paper, the Schwarz Information Criterion is used to select the appropriate lag length of the explanatory variables. Our analysis indicates that p, q and r are equal to 1. Furthermore, given the possible correlation between the lagged variables and the error terms, an instrumental variable estimator is required in order to have an unbiased estimate of Eqs. (2a)–(2c).

The estimation method widely applied in comparable studies in different fields of research (Bashiri Behmiri & Manso, 2012; Costantini & Martini, 2010; Jaunky, 2012a,b) is the system Generalized Method of Moments (GMM) proposed by Arellano and Bover (1995) and Blundell and Bond (1998). Following Holtz-Eakin et al. (1988), Arellano and Bond (1991) developed a GMM estimator that instruments the differenced variables that are not strictly exogenous with all their available lags in levels. Arellano and Bond also developed an appropriate test for autocorrelation, which can render some lags invalid as instruments. A problem with the original Arellano–Bond estimator is that lagged levels are poor instruments for first-differences if the variables are close to a random walk. Arellano and Bover (1995) describe how, if the original equation in levels is added to the system, additional instruments can be applied to increase efficiency. In this equation, variables in levels are instrumented with suitable lags of their own first-differences. We follow the instructions provided by Roodman (2006, 2009) in order to assess the validity of the instruments and of the instruments subsets, i.e. a high p-value of the Hansen J statistic (of at least 0.25) and of the difference-in-Hansen statistic. Furthermore, instruments are collapsed to limit instrument proliferation[11]. Finally, to control cross-sectional dependence, time dummies are included in the estimates.

In line with literature, three levels of causality are studied by means of a Wald test: short-run, long-run and strong causality. Considering first Eq. (2a), the "short-run Granger causality test" assesses the validity of the null hypothesis $H_0 : \eta_{im}^X = 0$ and

[9] In this paper the MG estimator proposed by Pesaran and Smith is preferred to the more commonly used dynamic OLS (DOLS) or to the fully modified OLS (FMOLS) procedures proposed by Saikkonen (1991) and Pedroni (2000) respectively, since this estimator is more appropriate to our datasets characterized by moderate length ($T = 20$).

[10] Our estimates are obtained by using the Stata routine proposed by Eberhardt (2012).

[11] More information about our estimations, such as the number of selected instruments in each equation and the usual diagnostic statistics, are available on request from the authors. However, in line with literature, the maximum number of instruments used is well below the threshold suggested by Roodman (2006, 2009) (maximum number of instruments = N, i.e. 37 in our case).

Table 4
Pedroni and Kao panel cointegration tests.

	China	France	Germany	Hong Kong	India	Indonesia	Italy	Netherlands	Spain	Turkey	UK	USA
Pedroni test												
Panel ν-statistic	−225.52 (1.00)	−292.31 (1.00)	−170.15 (1.00)	−233.59 (1.00)	−111.33 (1.00)	−74.01 (1.00)	−265.07 (1.00)	0.66 (0.25)	−40.92 (1.00)	−59.81 (1.00)	−368.05 (1.00)	−273.09 (0.06)
Panel ρ-statistic	−1.08 (0.14)	−0.68 (0.25)	−1.33 (0.09)	0.00 (0.50)	−0.11 (0.46)	−0.67 (0.25)	−1.04 (0.15)	−0.46 (0.32)	0.16 (0.56)	0.45 (0.67)	−0.29 (0.39)	−1.69 (0.04)
Panel *pp*-statistic	−2.92 (0.00)	−2.66 (0.00)	−2.09 (0.02)	−1.51 (0.06)	−1.62 (0.05)	−2.97 (0.00)	−4.15 (0.00)	−2.26 (0.01)	−2.45 (0.00)	−1.35 (0.09)	−2.03 (0.02)	−6.68 (0.00)
Panel adf-statistic	−4.96 (0.00)	−4.94 (0.00)	−1.55 (0.06)	−4.80 (0.00)	−2.30 (0.01)	−4.85 (0.00)	−3.36 (0.00)	−1.75 (0.04)	−2.84 (0.00)	−1.52 (0.06)	−3.75 (0.00)	−3.81 (0.00)
Group ρ-statistic	0.69 (0.75)	1.59 (0.94)	−1.23 (0.11)	2.21 (0.98)	1.24 (0.89)	1.58 (0.94)	0.62 (0.73)	1.32 (0.91)	1.48 (0.93)	2.55 (0.99)	0.49 (0.69)	0.18 (0.57)
Group *pp*-statistic	−2.69 (0.00)	−3.18 (0.00)	−6.59 (0.00)	−0.78 (0.22)	−2.06 (0.02)	−3.31 (0.00)	−4.08 (0.00)	−2.68 (0.00)	−2.54 (0.00)	−0.73 (0.23)	−3.31 (0.00)	−6.89 (0.00)
Group adf-statistic	−6.08 (0.00)	−6.89 (0.00)	−6.07 (0.00)	−4.47 (0.00)	−2.86 (0.00)	−3.98 (0.00)	−4.29 (0.00)	−2.04 (0.02)	−3.96 (0.00)	−1.64 (0.04)	−5.48 (0.00)	−3.31 (0.00)
Kao test												
ADF	−1.65 (0.05)	−4.11 (0.00)	−1.03 (0.15)	−1.77◆ (0.04)	−1.09◆ (0.14)	−2.36 (0.00)	−1.44◆ (0.07)	−3.21 (0.00)	−3.35 (0.00)	−5.43 (0.00)	−3.72 (0.00)	−3.82 (0.00)

Source: Authors' calculations on UN Comtrade data. *Notes:* The panel statistics are the within-dimension statistics while the group statistics are the between-dimension ones; the null hypothesis is no cointegration; *p*-values in parentheses; the lag length selection is based on SIC unless specified by adding a ◆ which indicates a user-specified lag length equal to 1; trend and intercept options are: "No deterministic trend" for all economies with the exception of the Netherlands, Spain and Turkey for which the option used was "No deterministic intercept or trend".

Table 5
Panel Granger causality tests.

		Short-run			Long-run ECT	Strong causality		
		$\Delta\ln X_{it}\Delta\ln X_{it}$	$\Delta\ln \mathrm{RP}_{it}$	$\Delta\ln \mathrm{GDPW}_t$		$\mathrm{ECT}/\Delta\ln X_{it}$	$\mathrm{ECT}/\Delta\ln \mathrm{RP}_{it}$	$\mathrm{ECT}/\Delta\ln \mathrm{GDPW}_t$
China	$\Delta\ln X_{it}$	–	5.12**	0.02	10.51***	–	7.08***	5.76***
	$\Delta\ln \mathrm{RP}_{it}$	1.09	–	0.27	0.80	0.84	–	0.53
France	$\Delta\ln X_{it}$	–	3.96*	1.77	7.02**	–	3.88**	3.71**
	$\Delta\ln \mathrm{RP}_{it}$	4.51**	–	3. 29*	4.10**	2.26	–	4.10**
Germany	$\Delta\ln X_{it}$	–	5.34**	4.76**	5.02**	–	3.02*	14.09***
	$\Delta\ln \mathrm{RP}_{it}$	1.54	–	0.26	1.60	0.82	–	0.84
Hong Kong	$\Delta\ln X_{it}$	–	0.04	0.4	5.11**	–	2.82*	2.55*
	$\Delta\ln \mathrm{RP}_{it}$	8.03***	–	1.22	3.93*	6.60***	–	5.51***
India	$\Delta\ln X_{it}$	–	0.02	8.06***	3.97*	–	3.03*	6.35***
	$\Delta\ln \mathrm{RP}_{it}$	5.32**	–	8.31**	11.19***	5.61***	–	7.69***
Indonesia	$\Delta\ln X_{it}$	–	5.87**	1.02	22.87***	–	25.43***	13.28***
	$\Delta\ln \mathrm{RP}_{it}$	1.00	–	3.87*	2.56	7.67***	–	1.96
Italy	$\Delta\ln X_{it}$	–	0.89**	0.5	30.25***	–	15.64***	15.39***
	$\Delta\ln \mathrm{RP}_{it}$	2.00	–	0.16	0.36	1.45	–	0.21
Netherlands	$\Delta\ln X_{it}$	–	0.34	0.47	4.08*	–	7.65***	2.63*
	$\Delta\ln \mathrm{RP}_{it}$	1.57	–	0.42	0.83	3.03*	–	0.42
Spain	$\Delta\ln X_{it}$	–	7.73***	0.17	10.94***	–	8.45***	11.54***
	$\Delta\ln \mathrm{RP}_{it}$	6.08**	–	0.71	2.51	3.66**	–	1.29
Turkey	$\Delta\ln X_{it}$	–	1.65	1.07	16.41***	–	12.11***	10.91***
	$\Delta\ln \mathrm{RP}_{it}$	0.00	–	8.60***	0.20	0.12	–	15.59***
UK	$\Delta\ln X_{it}$	–	2.25	4.35**	26.21***	–	13.14***	15.32***
	$\Delta\ln \mathrm{RP}_{it}$	8.16***	–	2.49	2.48	4.67***	–	3.14*
USA	$\Delta\ln X_{it}$	–	6.91**	0.85	32.28***	–	16.17***	16.16***
	$\Delta\ln \mathrm{RP}_{it}$	1.57	–	3.68*	0.23	0.82	–	2.40

Source: Authors' calculations on UN Comtrade data. *Note:* $^{*}(^{**})[^{***}]$ Indicates significance at 10(5)[1] per cent level.

$H_0 : \mu_{in}^X = 0$ for all i, m and n. The "long-run Granger causality test" checks for the significance of the ECT coefficient; in this case, the null hypothesis is $H_0 : \omega_i^X = 0$ for all i. Finally, strong causality assumes that the null hypotheses $H_0 : \omega_i^X = \eta_{im}^X = 0$ and $H_0 : \omega_i^X = \mu_{in}^X = 0$ jointly hold for all cross-sections. Similar null hypotheses are tested for Eqs. (2b) and (2c).

Table 5 reports the results of the Wald tests on the coefficients[12]. There is evidence of long-run causality from relative prices and world income to volumes for all economies in the sample. On the contrary, long-run Granger causality from volumes and income to relative prices is verified only for France, Hong Kong and India.

Short-run causality results are somewhat mixed. Evidence of short-run causality from relative prices to volumes is found for China, France, Germany, Indonesia, Italy, Spain, and the USA. Volumes are found to Granger-cause relative prices in the short run for France, Hong Kong, Spain and the UK. World income Granger-causes volumes in the short run in the case of Germany, India and UK while it Granger-causes prices in France, India, Indonesia, Turkey and the USA.

Finally, we provide joint Wald *F*-statistics for the interactive terms, i.e. the ECT and the explanatory variables, which give an indication of which variables bear the burden of short-run adjustment to re-establish long-run equilibrium, given a shock to the system (Asafu-Adjaye, 2000). As for Eq. (2a), both relative prices and world income are significantly strongly causal in re-adjusting towards equilibrium for all nations. As for Eq. (2b), results are mixed for strong causality as well. No evidence of causality is found for China, Germany, Italy, and the USA; both variables are responsible for adjustment in the case of Hong Kong, India and UK; world income is the only significant variable as for France and Turkey; relative prices bear the burden of short-run adjustment in Indonesia, the Netherlands and Spain.

Our analysis thus confirms that relative prices ($\ln \mathrm{RP}_{it}$) and world income ($\ln \mathrm{GDP}_t$) Granger-cause export volumes ($\ln X_{it}$) in the long-run, which is the time horizon we are interested in. Hence we can proceed to compare the estimated long-run price and income elasticities of the selected twelve top exporters.

4. Long-run export elasticities and their stability over time

4.1. Long-run export elasticities estimates

Once the direction of causality from prices and income to export volumes is established, we proceed with the discussion of the estimation results of Eq. (1). Our interest is focused on the coefficients β_i and γ_i, which indicate the export price and

[12] Since we are interested in the direction of causality between $\ln X_{it}$ and $\ln \mathrm{RP}_{it}$ (and $\ln \mathrm{GDPW}_t$) in order to understand whether relative prices (and income) determine export volumes or vice versa, we only report the results obtained for Eqs. (2a) and (2b). Estimates of Eq. (2c) were carried out but in general the diagnostic statistics did not allow us to draw any conclusions about causality, because at least one of the econometric requirements noted by Roodman (2006) was violated. This result is also in line with a priori economic principles where GDP is an exogenous variable, not dependent on relative prices and export volumes of any specific industrial sector.

income elasticity for each economy. The figures reported in Table 6 are the unweighted averages of the estimated coefficients across groups. All parameters are statistically significant at the 1% or 5% confidence level, with the only exceptions being price elasticity for Indonesia and income elasticity for the UK.

As regards price elasticities, they appear to be independent of AUVs, market shares and the development stage of an economy, i.e. whether it is advanced or emerging. Most values are within the range (−0.63, −0.84) with the exceptions of the USA, presenting the highest elasticity (−1.23), and Turkey and Spain, showing the lowest (−0.36 and −0.41, respectively).

As regards income elasticities (Table 6), China and Spain present the highest values (2.87 and 3.31, respectively), confirming the high growth rates recorded by their clothing exports during the two decades. India, Turkey, Indonesia, Germany and France show an estimated coefficient above 1 (only slightly for France), which means that their clothing exports grew at a faster rate than world income. The Netherlands and especially Italy show opposite trends. Finally, Hong Kong and the USA exhibit negative income elasticities as a consequence of the big fall in their clothing export volumes, probably due to a shift in their specialization towards more high-tech goods and services as well as their changing role in the global supply chain.

These results could be due to the fact that the MG estimator gives the same weight to all goods independently of their weight in trade volumes. For this reason, weighted coefficients for price and income elasticities are computed using as a weight the average share of export volume of each good over total clothing exports. The results are shown in Table 7.

It is clear that the new coefficients are not significantly different from the previous ones, although China and Hong Kong's price elasticities move closer to 1.

4.2. The stability of export elasticities

We now move on to investigate the stability of price and income elasticities over the time period considered. Rolling regressions are widely used in empirical papers on international trade, especially when economies, like those studied in this work, are characterized by deep structural changes, such as privatizations in the export sector, trade liberalization and increasing export shares coupled with relative high trade elasticities (see, among others, Aziz & Li, 2008; Basile, de Nardis, & Girardi, 2009). The technique is usually applied to check the stability of the long-run relationships between the variables of interest. If the estimated rolling period parameters prove to be in line with the long-run estimates shown in Table 6, the long-run relationships between the variables under investigation are confirmed. Otherwise, results should be treated with caution and policy-makers would be well advised to pay more attention to the dynamics of rolling-window parameters when evaluating the response of their economies to external shocks or designing industrial strategies. From this perspective, rolling-period elasticities are estimated using Eq. (1) and considering 12 different windows, starting from the time span 1992–2000 to 2003–2011 ($T = 9$ in each of them)[13].

Our regressions indicate that income elasticities are hardly stable in the sample economies (see Fig. 2). They generally show a decreasing trend, which is particularly marked for India and Indonesia from the late 1990s onwards and in the USA from the early 1990s. China exhibits the highest income elasticity in our sample (about 2.6). It is worth noticing that in this case the first period considered for estimation is the extended one 1992–2006, since rolling elasticities are not significant until the seventh window. The only countries showing an increasing trend in income elasticities are Germany and the Netherlands.

With regard to price elasticities (see Fig. 3), rolling estimates confirm the stability of the long run parameters reported in Table 6, with the exception of the Asian exporters (China, Hong Kong, India and Turkey). This finding is particularly interesting since, as noted in Section 2.1, it is clearly related to the big economic and institutional changes that have characterized these economies in the last twenty years (See Aziz & Li, 2008).

In particular, while long-run elasticities, as shown in Table 6, are in line with those of most advanced economies, the last rolling-window estimates show much higher coefficients, in absolute terms around 2.0, 1.6 and 1.2 for China[14], India and Hong Kong, respectively. Although these outcomes should be treated with caution because of the short time-window considered, they suggest that economies characterized by low/medium AUVs may exhibit a higher export price elasticity (for details about AUVs see Table 2). In fact, if we assume that AUVs are a good proxy for prices and that prices are a good proxy for quality, an economy producing higher-quality goods should be able to set higher export prices and profit from a more inelastic foreign demand. In a situation where there is high competition in the world market, the export price elasticity of economies producing higher-quality and hence higher-price goods should be lower than that of those producing lower-quality and lower-price goods. Hence we should observe an inverse relationship between the export price elasticities and the export prices of our sample economies (Baiardi, Bianchi, & Lorenzini, 2015). It is also worth noting that the increase in price elasticities occurs in about the same period as major changes in trade policy. This implies that liberalization has probably

[13] This technique is alternative to other popular methodologies testing panel structural breaks based on unit root tests (See, for example, Jewell, Lee, Tieslau, & Strazicich, 2003; Nag & Mukherjee, 2012). It is worth noticing that in this case it is extremely difficult to check for the numerous possible combinations of heterogeneous structural breaks that might occur (for details, see Jewell et al., 2003). Rolling regressions, on the other hand, have the advantage of identifying the parameters interested by possible time breaks and describing their dynamics over time. This analysis is thus particularly important in terms of economic policy.

[14] This result is in line with Aziz and Li (2008), who find a time-increasing price elasticity of Chinese total exports.

Table 6
Long-run estimates of Eq. (1).

	China	France	Germany	Hong Kong	India	Indonesia	Italy	Netherlands	Spain	Turkey	UK	USA
lnRP_{it}	-0.83^{***} (0.17)	-0.66^{***} (0.08)	-0.71^{***} (0.08)	-0.84^{***} (0.17)	-0.63^{***} (0.14)	-0.17 (0.10)	-0.76^{***} (0.12)	-0.79^{***} (0.11)	-0.41^{***} (0.13)	-0.36^{**} (0.15)	-0.85^{***} (0.07)	-1.23^{***} (0.18)
ln $GDPW_t$	2.87^{***} (0.28)	1.06^{***} (0.17)	1.33^{***} (0.25)	-0.60^{**} (0.27)	1.71^{***} (0.25)	1.41^{***} (0.25)	0.65^{***} (0.15)	0.93^{***} (0.25)	3.31^{***} (0.32)	1.52^{***} (0.25)	0.21 (0.16)	-1.26^{***} (0.31)
Constant	-13.16^{***} (3.02)	3.83^{**} (1.76)	1.30 (2.63)	22.67^{***} (2.86)	-3.53 (2.69)	0.29 (2.72)	8.89^{***} (1.64)	-4.69^{*} (2.65)	-20.96^{***} (3.37)	-0.99 (2.74)	12.71^{***} (1.70)	28.18^{***} (3.28)

Source: Authors' calculations on UN Comtrade data. *Note:* Standard errors in parentheses; $^{*}(^{**})[^{***}]$ indicates significance at 10(5)[1] per cent level.

Table 7
Long-run weighted estimates of Eq. (1).

	China	France	Germany	Hong Kong	India	Indonesia	Italy	Netherlands	Spain	Turkey	UK	USA
ln RP_{it}	-0.92^{***} (0.16)	-0.73^{***} (0.09)	-0.75^{***} (0.08)	-0.96^{***} (0.13)	-0.62^{***} (0.14)	-0.14 (0.11)	-0.74^{***} (0.12)	-0.72^{***} (0.13)	-0.43^{***} (0.14)	-0.39^{**} (0.15)	-0.87^{***} (0.05)	-1.25^{***} (0.18)
ln $GDPW_t$	2.69^{***} (0.30)	0.91^{***} (0.17)	1.24^{***} (0.27)	-0.50^{*} (0.27)	1.68^{***} (0.25)	1.36^{***} (0.24)	0.64^{***} (0.15)	0.86^{***} (0.27)	3.48^{***} (0.34)	1.53^{***} (0.25)	0.06 (0.15)	-1.32^{***} (0.31)

Source: Authors' calculation on UN Comtrade data. *Note:* Standard errors in parentheses; $^{*}(^{**})[^{***}]$ indicates significance at 10(5)[1] per cent level.

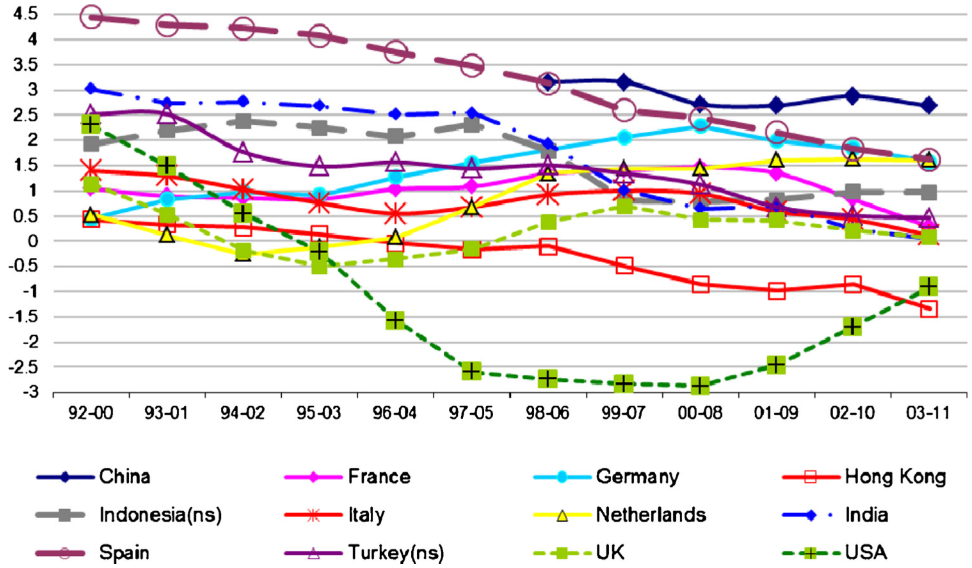

Fig. 2. Income elasticities estimated from rolling equations. *Source:* Authors' calculations on UN Comtrade data. *Notes:* With regard to China, the first period considered for estimation is the extended one 1992–2006 since rolling-period price elasticities are not significant until the 7th period.

contributed to making world demand more responsive to relative prices, especially for economies exporting low–medium price goods.

On the other hand, medium/high-price economies show fairly stable elasticities, below one in absolute terms. The liberalization process seems not to have impacted on the price elasticities of this cluster. This is probably because these economies specialize in higher quality goods as reflected in their higher AUVs. This implies a more inelastic world demand less influenced by price changes. Moreover, Spain's elasticity is the lowest for the whole group in the whole period, and Spain is the only country in the medium/high-price group to record a growth in the market share in both volumes and values. This suggests the emergence of a competitive gain mainly based on an improvement in quality and reputation. Indonesia and

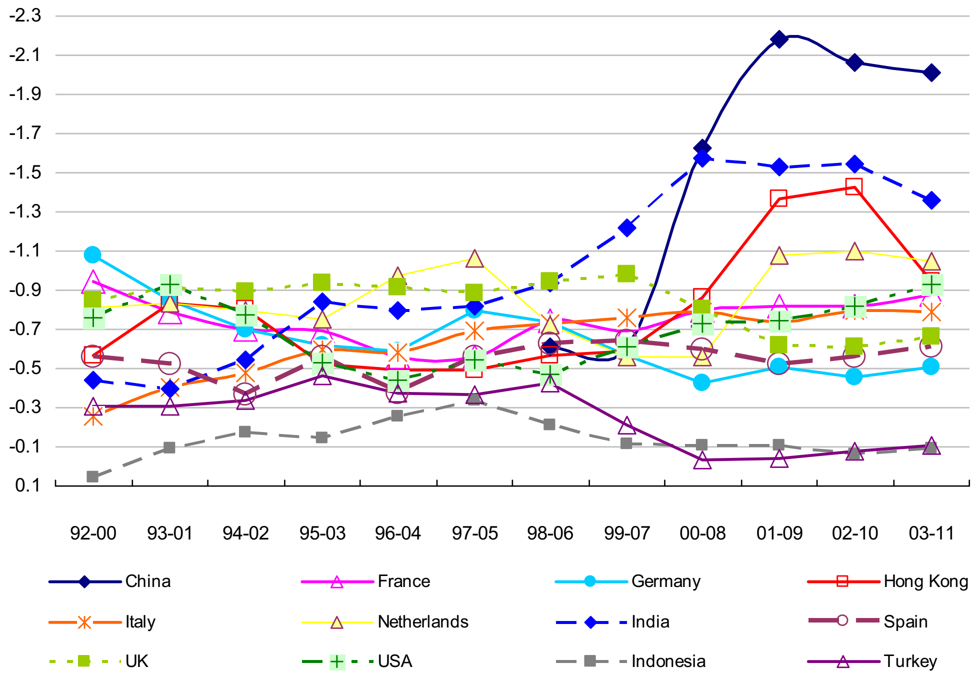

Fig. 3. Price elasticities estimated from rolling equations. *Source:* Authors' calculations on UN Comtrade data. *Notes:* With regard to China, the first period considered for estimation is the extended one 1992–2006 since rolling-period price elasticities are not significant until the 7th period.

Turkey, contrary to a priori expectations, show the rather puzzling feature of having the lowest elasticities in the sample in spite of being low-price countries. Further investigation of these features appears to be worthwhile.

5. Price elasticities, AUVs, export performance and perspectives

Various considerations apply to different economies regarding the characteristics, performance and prospects for clothing exports and production. First, China displays the highest price elasticity, which more than triples in the most recent sub-periods, and the second highest income elasticity after Spain. The high income elasticity indicates the persisting competitive strength of its clothing exports, but the large increase in the price elasticity signals that small changes in relative prices may have a big impact on export volumes. With regard to the other Asian exporters, Hong Kong and India's price elasticities show the same trend, although less pronounced, but unlike China, their increase is accompanied by a large fall in income elasticities and market shares. More specifically, it is worth noting that, around the end of the last century, Hong Kong modified its industrial organization from direct production to regional trading becoming a sourcing hub (Van Grunsven & Smakman, 2001). In the case of India, on the other hand, this is probably the consequence of a structural change in product specialization towards other sectors. Furthermore, Indonesia and Turkey show the rather puzzling feature that, contrary to a priori expectations, they exhibit the lowest elasticities in the sample in spite of being low-price countries. Possible explanations for this evidence found in the literature are competence-related quality and lead times for Turkey (Neidik & Gereffi, 2006; Tokatli & Kızılgün, 2009) and cheap but experienced labour force and industrial upgrading of production towards higher value products for Indonesia (Hassler, 2004).

The story is different for the advanced economies in our sample. The USA (Bair & Gereffi, 2003; Evans & Smith, 2006) and the European countries (Gibbon, 2002; Palpacuer, 2006) have been seriously affected by changes in the international division of labour. Globalization now means that distribution is mainly carried out by firms located in developed economies (Gereffi, 1994), where retailers and branders maintain direct control over design, marketing and related activities, and fix product standards for manufacturers in less developed economies (Gereffi, 1999). More advanced economies also show a constant decrease in employment and value added in the clothing industry, where a sizeable share of export volumes could be classified as processing trade or re-exports. This however tends not to be taken into account by official statistics.

These phenomena may explain the convergence process between European and Asian AUVs highlighted in Figs. 4 and 5, which show for each sample economy the dynamics of the combination of AUV (horizontal axis) and price elasticity (vertical axis) in the initial and final sub-periods 1992–2000 and 2003–2011, respectively. The choice of these two time windows is significant in light of the fundamental economic and institutional changes in the clothing industry noted in Section 2. The size of the circles reflects the market share in volumes for each exporter. The graphs are divided into four quadrants, partitioned by the mean AUV on the horizontal axis and the mean elasticity on the vertical axis. Obviously, the partition lines change over time; this is particularly relevant for price elasticities, whose mean value doubles in the most recent sub-period. This outcome is mainly due to the behaviour of the elasticities of medium-price economies, and especially China and India as explained above.

Figs. 4 and 5 show clearly that some countries in the sample, (Germany and Spain), perform better than others (France, the Netherlands and the UK) in their new role of export hubs, in terms of price and income elasticities and market shares. More specifically, Germany and France show the highest AUVs in the first sub-period, but rather unexpectedly their export prices decrease in the most recent years. At the same time, their price elasticities fall significantly so that their position shifts from the upper-right to the lower-right quadrant. The UK moves from the higher to the lower-right quadrant because of its reduction in price elasticity. A similar trend, although less pronounced, characterizes Spain. Italy deserves special attention,

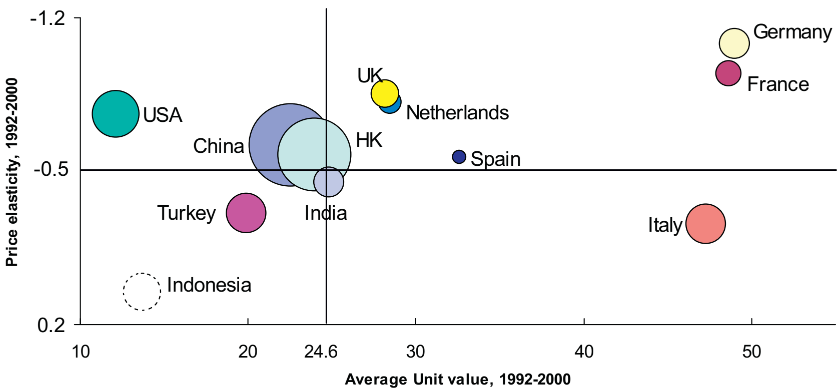

Fig. 4. Positioning of the various economies in terms of AUVs and price elasticities; time period 1992–2000. *Source:* Authors' calculations on UN Comtrade data. *Notes:* Price elasticities are estimated for the rolling period 1992–2000. The size of the circles is given by the average market share in volumes for the same period. In order to obtain a significant price elasticity, for China the considered sub-period is 1992–2006. The price elasticity for Indonesia is not significant for any sub-period.

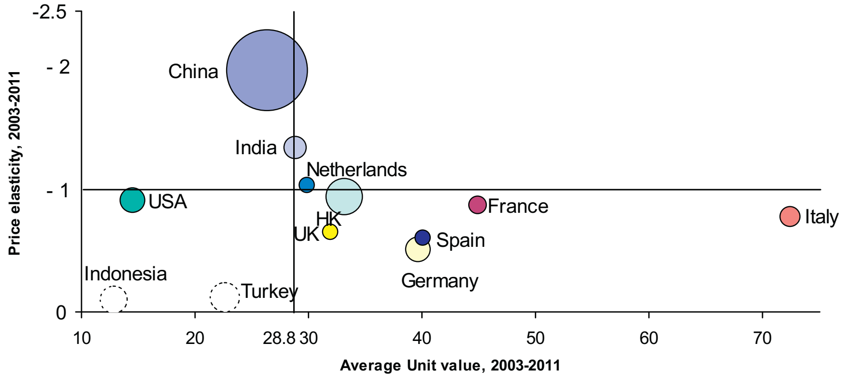

Fig. 5. Positioning of the various economies in terms of AUVs and price elasticities; time period 2003–2011. *Source:* Authors' calculations on UN Comtrade data. *Notes:* Price elasticities are estimated for the rolling period 2003–2011. The size of the circles is given by the average market share in volumes for the same period. The price elasticities for Indonesia and Turkey are not significant in the sub-period considered.

since it is the only advanced European country retaining a strong role for clothing in industrial production, with employment and value added figures much higher than those of other continental competitors. In Figs. 4 and 5, Italy lies in the quadrant of higher AUVs and more inelastic export demand in both sub-periods, and also differs from other exporters in its significant increase in AUVs during the sample period. In fact, Italy recorded an AUV of 92 USD in 2011, which is more than twice that of France, the country with the second highest AUV in the sample. This feature, however, is accompanied by a big fall in the estimated income elasticity and a rise in price elasticity.

There are various possible reasons for the high price of Italian products. On one hand, it could be a sign of quality-upgrading with a substantial selection process in the clothing industry that leaves only high-end firms to survive (for more details, Lissovolik, 2008). On the other hand, it could indicate a loss in competitiveness due to high input costs on the production side, and to a difficult absorption of high-price goods by international markets on the demand side. But despite the worrying constant decrease in Italy's export share, two elements deserve attention. First, the fact that the country's market share in values, although halved in twenty years, is still high and ranks third after China and Hong Kong. Second, the evidence shows that the demand for Italian goods continues to be inelastic (i.e. with a price elasticity below one in absolute values). In fact, it may be that an adjustment process is currently taking place, whereby Italy is positioning in the luxury goods segment, where there is market willingness to pay for high-price high-quality goods. Finally, as highlighted in the paper by Berthou and Emlinger (2010), it is also worth noting that high-quality goods, like those from Italy, are particularly susceptible to a fall in foreign imports during business-cycle recessions, as a consequence of a shift in consumer preference towards lower quality goods. This is a possible explanation for the Italian experience of the last decade, when, as the consequence of a long and deep financial crisis, German and Spanish products, notably characterized by lower unit values, appear to have substituted some Italian goods in international markets. This explanation is also consistent with our estimates of the dynamics of export income elasticities, which show a big decrease for Italy in the most recent sub-periods. A similar decrease can also be seen for France and the UK, and contrasts with the slight increase recorded by the Netherlands and the still high (although diminishing) values exhibited by Germany and Spain. Italy's lower income elasticity may also be due to the fact that exports are mainly directed towards slow-growing economies, with an accentuation of this phenomenon in recent years.

6. Conclusions

In this paper we estimate the price and income elasticities of 12 top exporters in the clothing industry for which data are available for the 20-year period 1992–2011.

First, we check whether the variables of interest in our estimates contain a unit root and are cointegrated. Second, we determine the pattern of causality by means of an Engle–Granger panel procedure. Once verified that relative prices and income Granger-cause export volumes for all economies, export function elasticities are estimated by means of a panel data approach. We find that most economies show similar long-run price elasticities, but their income elasticities differ considerably. Income elasticities are in fact very high (around 3) for China and Spain, and above 1 for India, Turkey, Indonesia, Germany and slightly above 1 for France, so that all these countries have a growth rate in clothing exports higher than that of world income. Moreover, looking at the stability of elasticities over time, rolling-period regressions indicate that income elasticities are generally decreasing, while price elasticities are more stable in time, with the single exception of the big Asian exporters in the sample, i.e. China, India and Hong Kong. This finding is in line with the fundamental economic and institutional structural changes that have characterized these economies in the last twenty years.

Different considerations apply to our clothing exporters regarding the characteristics and prospects of their foreign sales. Two main groups of economies can however be identified. On the one hand, the low-medium AUVs Asian exporters, i.e.

China, Hong Kong, India and Turkey, show unstable export price elasticities. This instability can be partly attributed to the liberalization of global trade and the end of the quota-system operated under the MFA. As noted in Section 2, trade liberalization from the 1990s to the early 2000s opened up advanced economies' markets to emerging economies' exports. In fact, it was precisely at the turn of the century that export price elasticities began to rise for China, India and Hong Kong. The liberalization of markets thus allowed economies with higher export price elasticities to gain market shares.

On the other hand, liberalization appears to have had no significant impact on the export price elasticities of the rest of the sample exporters. This is probably because their specialization model is based more on reputation than on price competitiveness. Moreover, the recent changes in the international division of labour have led to a shift of clothing production towards developing economies and a corresponding decrease in employment and value added in advanced economies. Although this is not captured by official statistics, a sizeable share of advanced economies' export volumes could be classified as processing trade or re-exports. Among advanced nations, Italy is the only exporter retaining a strong role for clothing in industrial production, with the highest AUV in the sample (92 USD in 2011, which is more than twice that of France, the second exporter with the highest AUV). There are various possible reasons for the high price of Italian products. On one hand, it could be a sign of quality-upgrading (Lissovolik, 2008) with a substantial selection process in the clothing industry that leaves only high-end firms to survive. On the other hand, it could indicate a loss in competitiveness due to high input costs on the production side, and to a difficult absorption of high-price goods by international markets on the demand side.

In sum, our analysis suggests that advanced economies, including Hong Kong but with the exception of Italy, specialized in the luxury goods segment, have changed their role in the clothing global value chain and moved towards an "organisational" position, whereby industrial production is located elsewhere and the economies become export hubs. In Asia, China confirms its leading position in clothing exports, as also revealed by its very high income elasticity. Finally, we highlight that the instability of price elasticities, together with their increasing trend for some Asian exporters, sounds a warning for future prospects.

Appendix A

Tables A1 and A2.

Table A1
Shares of clothing exports on total commodities exports, world and selected economies, beginning-of-period and end-of-period percentages.

	World	China	France	Germany	Hong Kong	India	Indonesia	Italy	Netherlands	Spain	Turkey	UK	USA
1992	3.68	19.67	2.27	1.94	16.77	15.02	9.32	6.86	1.93	1.10	28.40	1.92	0.94
2011	2.34	8.10	1.90	1.36	5.38	4.86	3.95	4.44	1.65	3.07	10.34	1.40	0.35

Source: Authors' calculation on WTO data.

Table A2
Clothing goods, SITC Classification Rev. 3.

Commodity code			Commodity description
84	**Clothing and accessories**		
	841	*Men, boys clothing, not knitted*	
		8411	Overcoats, outerwear, etc.
		8412	Suits and ensembles
		8413	Jackets and blazers, men's or boys', of textile materials, not knitted or crocheted
		8414	Trousers, bib and brace overalls, breeches and shorts, men's or boys'
		8415	Shirts
		8416	Underwear, nightwear, etc.
	842	*Women, girls clothing, not knitted*	
		8421	Overcoats, other coats, etc.
		8422	Suits and ensembles
		8423	Jackets and blazers, women's or girls', of textile materials, not knitted
		8424	Dresses, women's or girls', of textile materials, not knitted or crocheted
		8425	Skirts and divided skirts, women's or girls', of textile materials, not knitted
		8426	Trousers, bib and brace overalls, breeches and shorts, women's or girls'
		8427	Blouses, shirts
		8428	Underwear, nightwear, etc.
	843	*Men, boys clothing, knitted*	
		8431	Overcoats, car coats, capes, cloaks, anoraks (including ski jackets)
		8432	Suits, jackets, trousers, etc.
		8437	Shirts, knitted
		8438	Underwear, nightwear, etc.
	844	*Women, girls clothing, knitted*	
		8441	Overcoats, car coats, capes, cloaks, anoraks (including ski jackets)

Table A2 (*Continued*)

Commodity code		Commodity description
	8442	Suits, dresses skirts, etc.
	8447	Blouses, shirts and shirt blouses, women's or girls', knitted or crocheted
	8448	Underwear, nightwear, etc.
845	*Other textile apparel, nes*	
	8451	Babies' garments, clothes, accessories
	8452	Garment, felt, textile fabric
	8453	Jerseys, pullovers, cardigans, waistcoats and similar articles, knitted
	8454	Tshirts, singlets and other vests, knitted or crocheted
	8455	Brassieres, corsets, etc.
	8456	Swimwear
	8458	Other garments, not knitted
	8459	Other garments, knitted
846	*Clothing accessories, fabric*	
	8461	Accessories, not knitted
	8462	Hosiery, etc. knitted
	8469	Other madeup clothing accessories
848	*Clothing, nontextile; headgear*	
	8481	Leather apparel, accessories
	8482	Plastic, rubber apparel, etc.
	8483	Articles, accessories fur
	8484	Headgear, fittings, not else specified

Source: Authors' elaboration on the UN Comtrade database. *Note:* Only 4-digit level goods are used in the empirical analysis; the whole list is provided for completeness.

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