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Pricing to market of Italian exporting firms

by

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ABSTRACT

This paper investigates the pricing-to-market (PTM) behaviour of Italian exporting firms, using quarterly survey data by sector and by region over the period 1999q1-2005q2. A partial equilibrium imperfect competition model provides the structure according to which the orthogonality of structural shocks is derived. Impulse-response analysis shows non-negligible reactions of export-domestic price margins to unanticipated changes in cost competitiveness and in foreign and domestic demand levels, even though these effects appear to be of a transitory nature. For the period 1999-2001 a typical PTM behaviour emerges, while during the most recent years favourable foreign demand conditions allowed firms to increase their export-domestic price margins in face of a strong deterioration of their cost competitiveness. Macroeconomic implications of the observed PTM behaviour are also discussed.

Keywords: Pricing to market, survey data, panel-VAR models.

JEL Classification: E30, F31, F41.

NON-TECHNICAL SUMMARY

A large body of empirical studies shows that price differentiation across destination markets is the rule, rather than the exception, in the pricing behaviour of exporters of industrial countries. Exchange rate swings induce firms, endowed with some degree of market power, to adopt optimal (short run) price strategies, consisting in pricing the same good differently according to the markets where it is sold (pricing-to-market, PTM). Relevant macroeconomic consequences of PTM are incomplete exchange rate pass-through (ERPT) and deviations from the law of one price (LOP), such that the classical textbook reaction of the current account to exchange rate modifications (based on the Marshall-Lerner requirements on export and import price elasticities) can be considerably diluted. Necessary conditions for price discrimination, consequent on exchange rate movements, involve market-structure characteristics, functional forms of demand faced by firms in the various destination markets and degree of integration of trading countries.

Empirical work has shown that PTM and incomplete ERPT may be detected in a variety of industries and countries and, particularly, that these phenomena regard the behaviour of exporters of both large economies (like the US, Japan and Germany) and small countries (like Norway, Sweden and Korea).

In this paper we concentrate on Italy. Particularly, we look at the pricing behaviour of a sample of Italian exporting industrial firms, starting from disaggregated information at the sectoral/regional level, during the recent critical period, since 1999, when a significant reduction of the aggregate Italian market share in volume terms took place.

This paper aims at contributing to the empirical literature on PTM by developing a framework, which differs from previous works in several aspects. A *first* innovation concerns the data we use to test PTM of Italian firms. It is well known indeed that export unit value – the standard variable adopted in works dealing with pricing behaviour in foreign markets – provides a poor measure of export price, as it is affected by composition effects that blur the true meaning of its movements. We circumvent this problem by resorting to survey (qualitative) information where interviewed firms answer a precise question on the pricing policy they adopt when selling their product domestically and abroad: this is a unique piece of information which cannot be traced in any other statistical source. Furthermore, qualitative surveys on export activity of industrial firms provide a wide variety of indications on the behaviour of exporters (e.g., types of obstacles restraining exports, cost competitiveness, demand conditions at home and abroad) with the characteristics of being internally consistent (it is the

“same” individual firm providing all the requested information on its exporting activity). *Second*, a partial equilibrium model of imperfect competition provides the theoretical framework of reference for the empirical analysis. It presents the major advantage to provide empirical testable predictions on the key variable measuring the degree of PTM behaviour (i.e. the export-domestic price margin given by the ratio of the price of a good set by a firm for sale abroad relative to the price of the same good set for sale at home, converted to a common currency), which is indeed properly captured by a section of the survey used. Moreover, explicit conditions under which an exchange rate change translates into a rise or a fall in the export-domestic price margin are discussed. *Third*, in order to deal with the above-mentioned “intrinsic” endogeneity of the variables involved, the empirical analysis is based on the panel Vector AutoRegression (VAR) methodology. Such a technique seems to be particularly well suited for our purposes, since it applies the traditional VAR approach, where all variables in the system are treated as endogenous, within a panel data framework, which allows for unobserved individual heterogeneity.

Estimations results as well as dynamic simulation exercises suggest that while cost competitiveness factors have represented the major driving force for the sample of Italian firms in setting export prices over the period from 1999 to 2001, demand conditions in the domestic and in the foreign markets have exerted a relevant role in the most recent years, leading to an anomalous behaviour of the export-domestic price margin which was increasing in a period of rising competitive pressures. Sector analysis indicates that this phenomenon was mainly related to traditional industries, that is those that suffered the most because of foreign competition. This evidence notwithstanding, the peculiar pricing policy of Italian exporting firms turned to be limited in timing and in intensity. These findings seem to contrast PTM-based argumentations in explaining long-run deviations from the LOP.

“PRICING TO MARKET” DELLE IMPRESE ESPORTATRICI ITALIANE

SINTESI

Questo lavoro analizza le politiche di prezzo delle imprese esportatrici italiane, utilizzando dati trimestrali per settore e per ripartizione territoriale con riferimento al periodo 1999*q*1-2005*q*2. Un modello di equilibrio parziale di competizione imperfetta fornisce la struttura in base alla quale viene derivata l’ortogonalità degli *shock* strutturali. I risultati dell’analisi delle funzioni di risposta agli impulsi sembrano indicare che il margine di prezzo estero/interno risponda in misura statisticamente significativa a variazioni non attese di competitività di costo e nei livelli di domanda estera ed interna, sebbene tali effetti siano di natura transitoria. Per il periodo 1999-2001, vi è evidenza di un comportamento riconducibile al cosiddetto *pricing-to-market*, mentre negli anni più recenti condizioni particolarmente favorevoli sui mercati esteri hanno permesso alle imprese esportatrici di incrementare il margine di prezzo estero/interno nonostante il forte deterioramento della loro competitività di prezzo. Da ultimo sono discusse le implicazioni di carattere macroeconomico derivanti dal fenomeno del *pricing-to-market* empiricamente osservato.

Parole chiave: Pricing to market, data qualitativi, modelli VAR per dati longitudinali.

Classificazione JEL: E30, F31, F41.

CONTENTS

1 INTRODUCTION

A large body of empirical studies shows that price differentiation across destination markets is the rule, rather than the exception, in the pricing behaviour of exporters of industrial countries.[1] Exchange rate swings induce firms, endowed with some degree of market power, to adopt optimal (short run) price strategies, consisting in pricing the same good differently according to the markets where it is sold. Dornbusch (1987) investigates this phenomenon drawing on models of industrial organization, while Krugman (1987) formalizes it with the concept of pricing-to-market (PTM): if the domestic currency appreciates, a firm can react by reducing the domestic currency price of the good sold abroad to restrain the rise of the corresponding foreign price and the consequent fall of the (volume) market share. Motives behind the attempt at limiting the market share loss relate to the long-run investment made to establish in the market and the adjustment costs the firm has to incur when reducing volume of sales (these may be both reputation/brand-switching costs, as in Froot and Klemperer, 1989, and proper fixed costs in entering and exiting the market, as in Kasa, 1992). Relevant macroeconomic consequences of PTM are incomplete exchange rate pass-through (ERPT) and deviations from the law of one price (LOP), such that the classical textbook reaction of the current account to exchange rate modifications (based on the Marshall-Lerner requirements on export and import price elasticities) can be considerably diluted.

Irrespective of whether a general equilibrium framework (Betts and Devereux, 1996, among others) or a partial equilibrium model is concerned (Dornbusch, 1987; Krugman, 1987; Marston, 1990), necessary conditions for price discrimination, consequent on exchange rate movements, involve market-structure characteristics, functional forms of demand faced by firms in the various destination markets and degree of integration of trading countries. Particularly, a firm is able to vary its PTM if the following three conditions are met: *i)* imperfect competition, so that price does not equal marginal cost (or price elasticity of demand is not infinite); *ii)* variable price elasticity of demand in at least one destination market, so that PTM can vary in response to exchange rate changes; *iii)* segmented markets, so that product arbitrage by buyers and third parties is precluded (because resale of the product across nations is

1 For a comprehensive survey, see Goldberg and Knetter (1997).

costly, due to the existence of official and unofficial barriers).[2] It is worth noticing that while condition *ii)* is specific to PTM, conditions *i)* (imperfect competition) and *iii)* (market segmentation) are sufficient to allow price discrimination in circumstances other than those related to exchange rate changes.[3] Empirical work has shown that PTM and incomplete ERPT may be detected in a variety of industries and countries and, particularly, that these phenomena regard the behaviour of exporters of both large economies (like the US, Japan and Germany; see Knetter, 1989; Marston, 1990; Gagnon and Knetter, 1995; Mahdavi, 2002; Kikuchi and Sumner, 2002) and small countries (like Norway, Sweden and Korea; see Naug and Nyomen, 1996; Lee, 1997; and Adolfson, 1999). In this paper we concentrate on Italy. Particularly, we look at the pricing behaviour of a sample of Italian exporting industrial firms, starting from disaggregated information at the sectoral/regional level, during the recent critical period, since 1999, when a significant reduction of the aggregate Italian market share in volume terms took place.[4]

This paper aims at contributing to the empirical literature on PTM by developing a framework which differs from previous works in several aspects. A *first* innovation concerns the data we use to test PTM of Italian firms. Since Kravis and Lipsey (1971), it is well known, among analysts of international trade, that export unit value – the standard variable adopted in works dealing with pricing behaviour in foreign markets – provides a poor measure of export price, as it is affected by composition effects that blur the true meaning of its movements. We circumvent this problem by resorting to survey (qualitative) information where interviewed firms answer a precise question on the pricing policy they adopt when selling their product domestically and abroad: this is a unique piece of information which cannot be traced in any other statistical source. Furthermore, qualitative surveys on export activity of industrial firms

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2 Currency denomination of export prices, in the presence of nominal rigidities, is another potential source of observed PTM and of violations of the LOP. If firms set in advance (staggering) export prices in foreign currency, differentiation of selling prices to different markets may be the result of exchange rate surprises, rather than the deliberate choice of an optimising behaviour (Giovannini, 1988). Devereux (1997) provides an alternative explanation based on the assumption of price stickiness in the local currency of the (foreign) buyer.

3 For instance, it is possible for a firm to set prices differently according to the different cyclical demand conditions in the destination markets if the firm has market power and products cannot be arbitraged across borders.

4 Recent papers exploring the pricing behaviour of Italian exporters are those by de Nardis and Pensa (2004) and by Bugamelli and Tedeschi (2005): the former aims at detecting the degree of market power of Italian producers of traditional goods in foreign markets, estimating residual demand elasticities on the grounds of the approach of Goldberg and Knetter (1999); the latter is focused on the pricing strategies of Italian firms (across markets and sectors) during the nineties.

provide a wide variety of indications on the behaviour of exporters (e.g., types of obstacles restraining exports, cost competitiveness, demand conditions at home and abroad) with the characteristics of being internally consistent (it is the "same" individual firm providing all the requested information on its exporting activity); we make use extensively of this set of information in testing PTM.[5] *Second*, a reinterpretation of the partial equilibrium model of imperfect competition developed by Marston (1990) provides the theoretical framework of reference for the empirical analysis. It presents the major advantage to provide empirical testable predictions on the key variable measuring the degree of PTM behaviour (i.e. the export-domestic price margin given by the ratio of the price of a good set by a firm for sale abroad relative to the price of the *same* good set for sale at home, converted to a common currency), which is indeed properly captured by a section of the survey used. Moreover, explicit conditions under which an exchange rate change translates into a rise or a fall in the export-domestic price margin are discussed. *Third*, in order to deal with the above-mentioned "intrinsic" endogeneity of the variables involved, the empirical analysis is based on the panel Vector AutoRegression (VAR) methodology (Love, 2001). Such quite a novel technique seems to be particularly well suited for our purposes, since it applies the traditional VAR approach, where all variables in the system are treated as endogenous, within a panel data framework, which allows for unobserved individual heterogeneity.

Estimations results as well as dynamic simulation exercises suggest that while cost competitiveness factors have represented the major driving force for the sample of Italian firms in setting export prices over the period from 1999 to 2001, demand conditions in the domestic and in the foreign markets have exerted a relevant role in the most recent years, leading to an anomalous behaviour of the export-domestic price margin which was increasing in a period of rising competitive pressures. Sector analysis indicates that this phenomenon was mainly related to traditional industries, that is those that suffered the most because of foreign competition. This evidence notwithstanding, the peculiar pricing policy of Italian exporting firms turned to be limited in timing and in intensity. These findings seem to contrast PTM-based argumentations in explaining long-run deviations from the LOP.

The paper is structured as follows. In section 2 we present the theoretical framework used to model PTM behaviour of exporting firms. Dataset characteristics are illustrated in section 3. Section 4 reports the empirical

[5] A former attempt to study export pricing policies of Italian firms using survey information is in Pupillo and Zimmermann (1991).

specification and estimation results. In section 5 we discuss dynamic simulation exercises. Section 6 concludes.

2 THEORETICAL FRAMEWORK

Formalization of the theoretical framework to test PTM of Italian exporting firms is based on the model originally proposed by Marston (1990). We consider a firm operating in sector *i* endowed, for whatever reason, with some degree of market power.[6] As in Giovannini (1988) and Marston (1990), the firm sells its product in two distinct markets, labelled respectively as Home (H) and Foreign (F) destinations. The two markets are segmented, so that the firm is potentially able to practice different prices in each of them. Moreover we assume, for the sake of simplicity, that the good sold to the two markets is produced in the Home country by a single-plant firm. This makes marginal costs independent of the destination markets; yet, marginal costs may respond to changes in total sales (caused by modifications of either domestic or foreign sales). The imperfectly competitive firm will set prices in each market in order to maximize its profit.

Let P_{it}^{j} be the domestic currency price practiced by the firm of sector i at time t in each market j ($j = H, F$), X_{it}^{j} be the volume of sales to each market, E_t the exchange rate of H 's currency per unit of F 's. We consider, moreover, a cost function, C, a price index of production inputs, W_{it}, the price of competitors expressed in the currency of market *F*, P_{it}^{f}, and the demand of each market for all possible varieties of the considered product, V_{it}^{j}. Profit maximization of firm *i* at each time *t*

$$\max_{X_{it}^{j}} \left[\left(\sum_j P_{it}^{j} \cdot X_{it}^{j} \right) - C \cdot \left(\sum_j X_{it}^{j}, W_{it} \right) \right] \tag{1}$$

with $$X_{it}^{j} = X_{it}^{j}(P_{it}^{j} / E_t ; P_{it}^{f} ; V_{it}^{j})$$

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6 It might derive from monopolistic competition and product differentiation or from any other market imperfection.

leads to the first-order conditions which define the price set in each destination market as a mark-up over marginal cost

$$P_{it}^{j} = MC_{it} \cdot \left(\frac{\eta_{it}^{j}}{\eta_{it}^{j} - 1} \right) \qquad (2)$$

where MC_{it} and $\eta_{it}^{j} = -\left(\partial X_{it}^{j} / \partial P_{it}^{j}\right) \cdot \left(P_{it}^{j} / X_{it}^{j}\right)$ indicate, respectively, the marginal cost and the price elasticity of demand faced by firm i in H and F markets.

Since the marginal cost is independent of j (destination markets), price differences in the two markets (and hence geographical differences of the mark-ups) depend only on the price elasticity of demand in the two destination markets. In turn, such elasticities reflect the characteristics (degree of convexity) of the demand schedules. Models of price discrimination leading to PTM hence require particular functional forms of the demand curves, such that the price elasticity of the demand schedule, η_{it}^{j}, is not constant in at least one destination market.[7] In general, demand functions must be less convex (or more linear) than the constant price elasticity demand schedule in order to have positive PTM elasticities (with respect to the exchange rate). Under this condition, an appreciation of the domestic currency leads to an increase of the foreign currency price practiced in the export market less than proportional to the exchange rate appreciation; with variable demand price elasticities,

[7] If both demand curves in the two markets have constant price elasticities – as in a log-linear demand schedule – then the mark-ups in each destination market are constant; in this case prices practiced in *H* and *F* are affected by a common marginal cost and the ratio between the price made in *F* and the price made in *H* (expressed in the same currency) is invariant to currency movements; this means that there may be a constant wedge between the two prices, but PTM does not change in response to variations of the exchange rate. Whether ERPT is complete or not depends, in this case, on the behaviour of the marginal cost: if the latter is increasing in output (total sales), ERPT is partial (i.e. an appreciation of the domestic currency, negatively affecting total sales, gives rise to a decrease of the marginal cost and, hence, to a reduction of both the domestic and foreign price expressed in the home currency). As a consequence, the export price expressed in the foreign currency increases less than proportionally to the exchange rate change (see the Appendix).

incomplete ERPT is always the rule whatever the behaviour of the marginal cost in response to total sales changes.[8]

Former considerations can be succinctly described as follows. Labelling by $R_{it} = P_{it}^F / P_{it}^H$ the export-domestic price margin (in the spirit of Marston, 1990), the elasticity $\varepsilon_{(P_{it}^F / P_{it}^H)E_t}$ of R_{it} with respect to exchange rate is given by

$$\varepsilon_{(P_{it}^F / P_{it}^H)E_t} = \frac{\partial R_{it}}{\partial E_t} \cdot \frac{E_t}{R_{it}} = \varepsilon_{P_{it}^F E_t} - \varepsilon_{P_{it}^H E_t} \qquad (3)$$

where both terms $\varepsilon_{P_{it}^F E_t}$ and $\varepsilon_{P_{it}^H E_t}$ are positive, as shown in the Appendix. On the grounds of these signs, the elasticity of R_{it} with respect to the exchange rate is positive and less than 1 – such that an appreciation (depreciation) of the domestic currency leads to a less than proportional reduction of the export-domestic price margin (and vice versa) – if the overall reaction of P_{it}^F following an exchange rate movement is larger than the reaction of P_{it}^H. This is certainly the case when marginal cost affects domestic and foreign prices in the same way (see the Appendix).

In the empirical specification of the model we focus on the level of the export-domestic price margin, R_{it}. This implies that, while levels of common marginal costs wipe out, R_{it} will depend on the influences activated by both exchange rate changes and movements of all the other factors, which include, as described in (1), the price of competitors P_{it}^f, named in the F's currency, and the cyclical conditions of demand in the domestic and foreign markets, indicated with V_{it}^j.

To make explicit the model for empirical testing, we take the logarithmic transformation of (2), linearise the mark-up in the two markets around respective average values through a first order Taylor approximation and take the difference between the logarithms of prices in the Foreign and Home markets (see Adolfson, 1999). All this yields

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8 In the model by Caves and Jones (1977), the phenomenon of "dumping" in international trade arises when a *monopolistic* profit maximizing firm, facing a higher elasticity of demand abroad than at home, is able to discriminate between foreign and domestic markets. As a consequence, it will charge a lower price abroad than at home. Conversely, Brander and Krugman (1983) show how dumping may arise for "systematic" reasons associated with *oligopolistic* behaviour rather than "accidental" differences in country demands.

$$r_{it} \cong p_{it}^{F} - p_{it}^{H} = \beta_1 \cdot (e_t + p_{it}^{f}) + \beta_2 \cdot f_{it} + \beta_3 \cdot h_{it} \quad (4)$$

where $p_{it}^{F} = \ln P_{it}^{F}$, $p_{it}^{H} = \ln P_{it}^{H}$, $e_t = \ln E_t$, $p_{it}^{f} = \ln P_{it}^{f}$, $f_{it} = \ln V_{it}^{F}$, $h_{it} = \ln V_{it}^{H}$ and the constant, grouping the time independent terms (average values) that come from the linearisation, is omitted.

In (4), the PTM behaviour of firms is captured by the parameter β_1 which measures the response of the export-domestic price margin to changes in competitors' prices expressed in domestic currency. Since price setting, described in (4), bases export prices on the overall competitive conditions in foreign markets (rather than just the nominal exchange rate), this formulation allows to consider the possibility of PTM behaviour in markets (and with respect to competitors) with which exporting firms share fixed-exchange rate regimes or even the same currency.[9]

3 DATA

Data are taken from the quarterly ISAE (Institute for Economic Studies and Analyses) survey. It covers a representative sample of manufacturing exporting firms with more than 10 employees and operating in Italy. The sample (around 2,000 firms) is random and stratified according to the number of employees, the sector and the regional location of the firm. We base our analysis on quarterly data covering the period between January 1999 and June 2005. Survey data not only constitute a valuable complement to statistical sources, which involve typically quantitative variables, but also provide a reliable (and direct) picture of agents' opinions about the dynamics of variables influencing their position in the markets. Moreover, the use of survey data makes it possible to work with a panel dimension (information are set at region/sector and time level) which is precluded by other statistical sources.

[9] In the empirical specification, we hence follow Marston (1990) who links (permanent) changes in the export price setting to movements of the real (rather than nominal) exchange rate; in our formulation $e_t + p_{it}^{f}$ is the (logarithm of the) numerator of the real exchange rate specific to sector *i*.

Table 1 reports the distribution of sample firms by sector and by region.[10] The number of firms entering in our panel dataset represents more than four-fifth of the total (1,275 out of 1,525), excluding sector "DF-Manufacture of coke, refined petroleum products and nuclear fuel", which contains numerous missing values in all regions.

Tab. 1 **ISAE survey on exporting firms**

(a) Number of firms by sector and by region	North West	North East	Centre	Total by sector
DA - Food products, beverages, tobacco	24	40	15	79
DB - Textile, textile products	59	61	68	188
DC - Leather, leather products	7	25	43	75
DD - Wood, wood products	8	22	10	40
DE - Pulp, paper, paper product, publishing, printing	13	21	17	51
DG - Chemicals, chemical products, man-made fibres	25	15	21	61
DH - Rubbers and plastic products	32	27	17	76
DI - Other non-metallic mineral products	20	35	23	78
DJ - Basic metals, fabricated metal products	61	73	27	161
DK - Machinery and equipment n.e.c.	47	94	39	180
DL - Electrical and optical equipment	41	35	25	101
DM - Transport equipment	31	19	13	63
DN - Furniture, n.e.c.	18	64	40	122
Total by region	386	531	358	1275

(b) Definition of the regions	
North West	Piemonte, Lombardia, Liguria, Valle D'Aosta
North East	Friuli-Venezia Giulia, Veneto, Emilia Romagna, Trentino Alto-Adige
Centre	Lazio, Toscana, Marche, Umbria

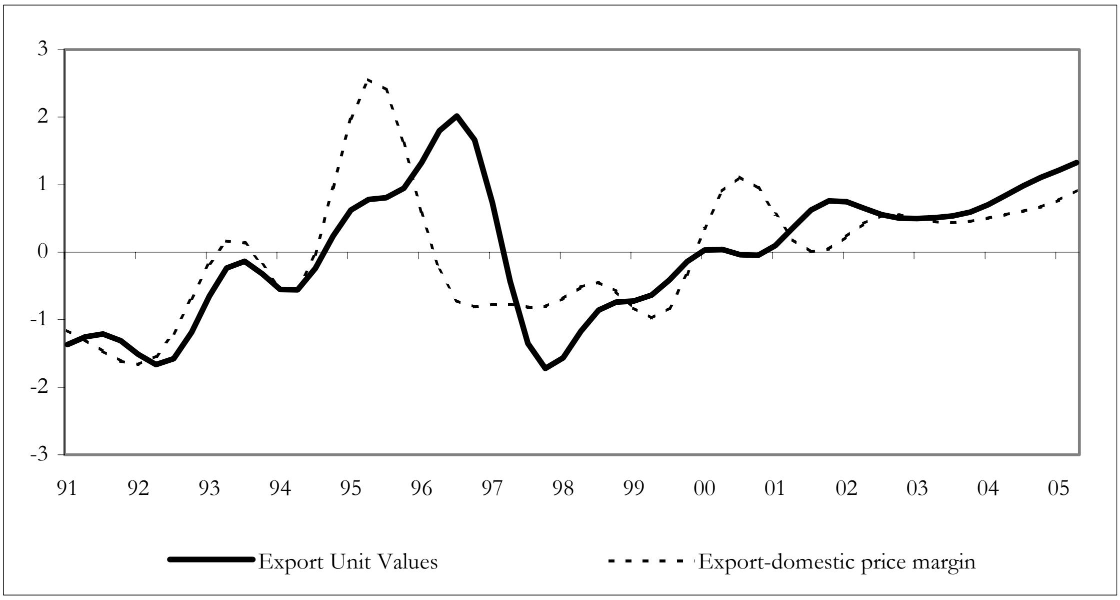
Note. Classification by sector according to NACE (rev. 1.1) - ATECO 2002.

Here is an overview of the variables involved in the empirical analysis. From the point of view of the PTM analysis, the most relevant question of the survey concerns the indication about the export-domestic price margin practiced by the firm. Firms are asked to indicate whether the prices charged in foreign markets are higher than, equal to, or lower than those practiced in the domestic market. Individual answers are aggregated by computing an indicator which takes values ranging between -100 and 100. In this work, relative prices are argued to fit the testing of the economic hypotheses discussed in Section 2 much better than any other information provided by traditional sources for several reasons. First, survey data on export-domestic price margins allow

[10] In the empirical investigation, firms belonging to the South of Italy (Sicilia, Sardegna, Calabria, Puglia, Basilicata, Molise, Campania, Abruzzo) are not included because of the scant relevance of several sectors in that region.

dealing with proper export price rather than export unit values (EUV), which are an inaccurate measure of export prices, being affected by important composition effects. Second, as stressed by Goldberg and Knetter (1997), indeed, it is important that the relevant variable in the analysis is the export/domestic price for the *same good* produced by the *same firm*, since it allows capturing price differences from two different transactions.[11] As illustrative example, Figure 1 shows a comparison of the EUV/producer prices ratio and its survey qualitative counterpart (P_{it}^{F} / P_{it}^{H}), over the period 1991*q*1-2005*q*2. Even though both series present similar behaviours over time (correlation equal to 0.63), export-domestic price margins seem to anticipate the quantitative series. Moreover, in the most recent years, the qualitative indicator exhibits a slightly higher volatility.

Fig. 1 Export unit values and export-domestic price margin Italy economy: 1991*q*1-2005*q*2.



Note. Cycle plus trend from Band-Pass filter on normalized data.

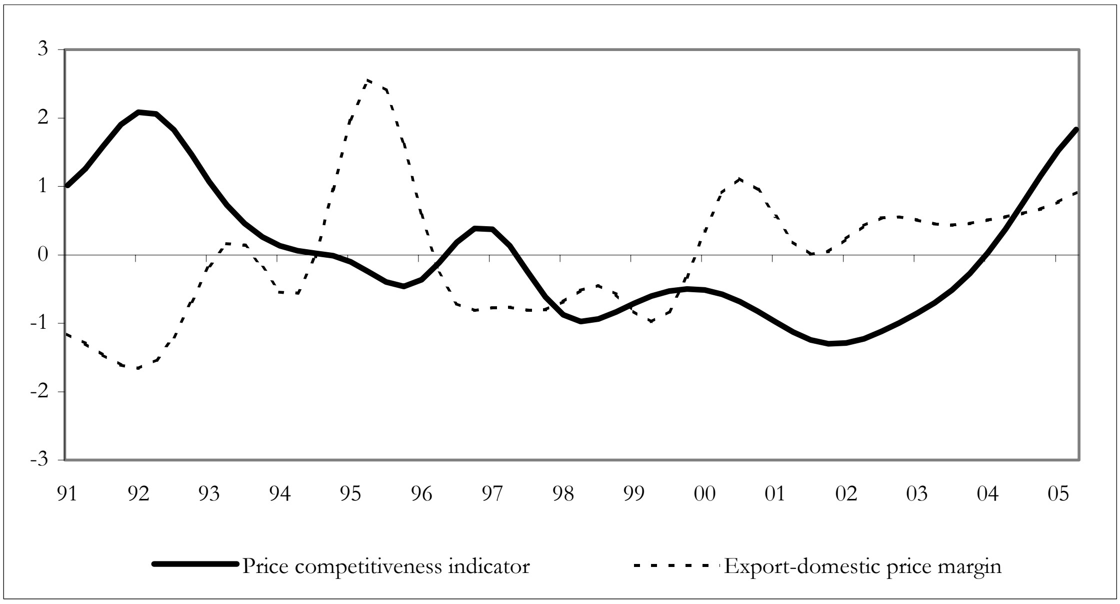
Other relevant information coming from the ISAE survey regard the domestic and foreign demand levels. Firms are asked to indicate whether the domestic (foreign) demand level increased, remained unchanged or decreased over the period of reference. The qualitative variable is constructed as the difference between the relative frequency of firms that declare an increase and

[11] Consider, for instance, Italian foot-wears. Comparing domestic prices to those charged abroad by the same “shoe-firm” makes undoubtedly more sense than comparing the price of foot-wears produced by different firms at home and abroad.

the relative frequency of those indicating a decline in the domestic (foreign) demand levels. Accordingly, the values of the two indicators range from -100 to 100. In terms of our theoretical model, domestic and foreign orders are used as proxies of the cyclical demand conditions in the two markets, which are the other factors affecting the behaviour of the export-domestic price margin.

The other qualitative variable we use refers to cost competitiveness obstacles. The price competitiveness indicator is represented by the share of firms that experienced impediments in foreign markets due to price competitiveness from foreign exporters. As a result, such a variable is bounded to belong to the interval between 0 and 100. In terms of the variables of the theoretical model, the price competitiveness indicator synthesizes the interaction between nominal effective exchange rate and the price pressure exerted by foreign exporters ($Q_{it} \equiv P_{it}^{f} \cdot E_{t}$). Figure 2 illustrates the evolution over time of export-domestic price margins and the price competitiveness indicator, Q_{it}, over the period 1991*q*1-2005*q*2 for the aggregate Italian manufacturing sector as a whole.

Fig. 2 Price competitiveness indicator and export-domestic price margin Italy economy: 1991*q*1-2005*q*2.



Note. Cycle plus trend from Band-Pass filter on normalized data.

A strong negative correlation (–0.37) between the two series characterizes the entire period considered, even though they co-move in the last four years (the correlation turns to be positive and equal to 0.86), when the perception of increasing competitive pressures in foreign markets was accompanied by a rise in the export-domestic price margin.

Finally, Table 2 presents some descriptive statistics of the qualitative series taken from the surveys. The indicators refer to the full post-euro period 1999*q*1-2005*q*2 (upper panel) as well as the two sub-samples 1999*q*1-2001*q*4 and 2002*q*1-2005*q*2 (central and lower panel).

Tab. 2 **Descriptive statistics**

Variable	N. Obs.	Mean	Std. Dev.	Min	Ma	Percentile		
						25	50	75
Period: 1999*q*1-2005*q*2								
Q	1014	18.23	14.32	0.00	80.03	7.05	16.48	26.16
V^F	1014	-17.80	18.18	-75.10	49.30	-30.60	-17.60	-5.48
V^D	1014	-15.93	16.99	-79.20	49.34	-27.88	-16.05	-4.00
P^F/P^H	1014	4.36	17.84	-82.46	75.03	-5.16	2.36	13.63
Period: 1999*q*1-2001*q*4								
Q	468	15.43	12.62	0.00	80.03	4.67	13.91	22.51
V^F	468	-12.20	17.92	-75.10	49.30	-24.11	-11.75	0.06
V^D	468	-11.16	18.54	-79.20	49.34	-23.60	-8.34	1.62
P^F/P^H	468	3.30	18.41	-82.46	69.28	-7.41	1.28	14.05
Period: 2002*q*1-2005*q*2								
Q	546	20.64	15.24	0.00	79.73	9.32	18.68	29.02
V^F	546	-22.60	17.01	-67.50	31.40	-35.00	-23.35	-11.36
V^D	546	-20.01	14.35	-73.20	27.70	-29.70	-20.35	-10.18
P^F/P^H	546	5.26	17.30	-70.23	75.03	-3.90	3.06	13.37

All of the series take values consistent with the range within which they are expected to fall. More interestingly, the mean value of export-domestic price margins is positive, suggesting the existence of some degree of market power and of market segmentation, despite the fact that a large share of exports was sold in EMU countries.

4 ECONOMETRIC ANALYSIS: EMPIRICAL SPECIFICATION AND ESTIMATION RESULTS

Survey data have been properly transformed in order to perform the econometric analysis. For those variables constructed as differences between positive and negative outcomes from multiple-choice-answers in the survey (P_{it}^F / P_{it}^H , V_{it}^F , V_{it}^H) the following transformation is used

$$-\ln\left[\frac{200}{Z_{it}+100}-1\right]$$

where Z_{it} is a generic series which represents P_{it}^F / P_{it}^H , V_{it}^F or V_{it}^H . The conversion allows to render unbounded otherwise limited variables. Conversely, the price competitiveness indicator, Q_{it} , is bounded within the interval [0,100] and transformed as

$$\ln\left[1+Q_{it}\right]$$

4.1 Empirical specification

For the empirical implementation of model (4), we consider the possibility that omitted variables may contribute to determine export-domestic price margins. Some of these variables may be constant over time but vary between sectors/regions, and others may be fixed between sectors/regions but vary over time. Secondly, we try to control for all potential endogeneity sources: simultaneity and measurement errors. We start from the following model

$$r_{it} = \gamma_1^R \cdot q_{it} + \gamma_2^R \cdot f_{it} + \gamma_3^R \cdot h_{it} + \alpha_i^R + \delta_t^R + \nu_{it}^R + m_{it}^R \tag{5}$$

$$v_{it}^R = \rho^R \cdot v_{i,t-1}^R + e_{it}^R \ , \ \left|\rho^R\right| < 1 \ , \ e_{it}^R, m_{it}^R \sim MA(0)$$

where α_i^R is a sector/region-specific intercept, δ_t^R is a time-specific intercept and v_{it}^R is a possibly autoregressive shock. We are interested in consistent estimation of the parameters ($\gamma_1^R, \gamma_2^R, \gamma_3^R, \rho^R$), maintaining that both the price competitiveness factor q_{it} and (foreign and domestic) demand levels are potentially correlated with the sector/region-specific effects (α_i^R), and with both sector/regional shocks (e_{it}^R) and serially uncorrelated measurement errors (m_{it}^R).

Price competitiveness factors are explained by no other variables, except for unobserved sector/region-specific and time effects

$$q_{it} = \alpha_i^Q + \delta_t^Q + \nu_{it}^Q + m_{it}^Q \tag{6}$$

$$v_{it}^Q = \rho^Q \cdot v_{i,t-1}^Q + e_{it}^Q \ , \ \left|\rho^Q\right| < 1 \ , \ e_{it}^Q, m_{it}^Q \sim MA(0)$$

while foreign demand depends on price competitiveness factors

$$f_{it} = \gamma_1^F \cdot q_{it} + \alpha_i^F + \delta_t^F + \nu_{it}^F + m_{it}^F \tag{7}$$

$$v_{it}^F = \rho^F \cdot v_{i,t-1}^F + e_{it}^F \ , \ \left|\rho^F\right| < 1 \ , \ e_{it}^F, m_{it}^F \sim MA(0)$$

Finally, domestic demand is influenced by current price competitiveness factors as well as foreign demand[12]

$$h_{it} = \gamma_1^H \cdot q_{it} + \gamma_2^H \cdot f_{it} + \alpha_i^H + \delta_t^H + \nu_{it}^H + m_{it}^H \tag{8}$$

$$v_{it}^H = \rho^H \cdot v_{i,t-1}^H + e_{it}^H \ , \ \left|\rho^H\right| < 1 \ , \ e_{it}^H, m_{it}^H \sim MA(0)$$

As usual, the structural model (5)-(8) admits the dynamic representation

$$R(L) \cdot r_{it} = \gamma_1^R \cdot R(L) \cdot q_{it} + \gamma_2^R \cdot R(L) \cdot f_{it} + \gamma_3^R \cdot R(L) \cdot h_{it} + \tilde{\alpha}_i^R + \tilde{\delta}_t^R + \tilde{m}_{it}^R$$

$$Q(L) \cdot q_{it} = \tilde{\alpha}_i^Q + \tilde{\delta}_t^Q + \tilde{m}_{it}^Q$$

[12] The idea that foreign development affects what happens within national boundaries (rather than the other way around) is consistent with the small open economy assumption for the Italian economic system.

$$F(L)\cdot f_{it} = \gamma_1^F \cdot F(L)\cdot q_{it} + \tilde{\alpha}_i^F + \tilde{\delta}_t^F + \tilde{m}_{it}^F$$

$$H(L)\cdot h_{it} = \gamma_1^H \cdot H(L)\cdot q_{it} + \gamma_2^H \cdot H(L)\cdot f_{it} + \tilde{\alpha}_i^H + \tilde{\delta}_t^H + \tilde{m}_{it}^H$$

where L denotes the lag operator and with $\varphi(L) = \left(1-\rho^{\varphi}L\right)$, $\tilde{\delta}_t^{\varphi} = \varphi(L)\cdot\delta_t^{\varphi}$, $\tilde{\alpha}_i^{\varphi} = \alpha_i \cdot \left(1-\rho^{\varphi}\right)$, $\tilde{m}_{it}^{\varphi} = e_{it}^{\varphi} + \varphi(L)\cdot m_{it}^{\varphi}$, $\varphi = R,Q,F,H$.

More compactly, the model can be written as follows

$$\underbrace{\begin{bmatrix} 1 & 0 & 0 & 0 \\ \gamma_1^F & 1 & 0 & 0 \\ \gamma_1^H & \gamma_2^H & 1 & 0 \\ \gamma_1^R & \gamma_2^R & \gamma_3^R & 1 \end{bmatrix}}_{\mathbf{A}_0} \cdot \underbrace{\begin{bmatrix} q_{it} \\ f_{it} \\ h_{it} \\ r_{it} \end{bmatrix}}_{\mathbf{y}_{it}} = \underbrace{\begin{bmatrix} \rho^Q & 0 & 0 & 0 \\ \gamma_1^F\rho^F & \rho^F & 0 & 0 \\ \gamma_1^H\rho^H & \gamma_2^H\rho^H & \rho^H & 0 \\ \gamma_1^R\rho^R & \gamma_2^R\rho^R & \gamma_3^R\rho^R & \rho^R \end{bmatrix}}_{\mathbf{A}_1} \cdot \underbrace{\begin{bmatrix} q_{i,t-1} \\ f_{i,t-1} \\ h_{i,t-1} \\ r_{i,t-1} \end{bmatrix}}_{\mathbf{y}_{i,t-1}} + \underbrace{\begin{bmatrix} \tilde{\alpha}_i^Q \\ \tilde{\alpha}_i^F \\ \tilde{\alpha}_i^H \\ \tilde{\alpha}_i^R \end{bmatrix}}_{\boldsymbol{\alpha}_i} + \underbrace{\begin{bmatrix} \tilde{\delta}_t^Q \\ \tilde{\delta}_t^F \\ \tilde{\delta}_t^H \\ \tilde{\delta}_t^R \end{bmatrix}}_{\boldsymbol{\delta}_t} + \underbrace{\begin{bmatrix} e_{it}^Q \\ e_{it}^F \\ e_{it}^H \\ e_{it}^R \end{bmatrix}}_{\mathbf{e}_{it}} \quad (9)$$

under the assumption that measurement errors exist but are constant over time.[13] Recasting the model in matrix form, we have

$$\mathbf{y}_{it} = \mathbf{A}_0^{-1}\cdot\mathbf{A}_1\cdot\mathbf{y}_{i,t-1} + \mathbf{A}_0^{-1}\cdot\boldsymbol{\alpha}_i + \mathbf{A}_0^{-1}\cdot\boldsymbol{\delta}_t + \mathbf{A}_0^{-1}\cdot\mathbf{e}_{it}$$

or

$$\mathbf{y}_{it} = \mathbf{\Pi}_1\cdot\mathbf{y}_{i,t-1} + \mathbf{a}_i + \mathbf{d}_t + \boldsymbol{\varepsilon}_{it} \quad (10)$$

where $\mathbf{\Pi}_1 = \mathbf{A}_0^{-1}\cdot\mathbf{A}_1$, $\mathbf{a}_i = \mathbf{A}_0^{-1}\cdot\boldsymbol{\alpha}_i$, $\mathbf{d}_t = \mathbf{A}_0^{-1}\cdot\boldsymbol{\delta}_t$ and $\boldsymbol{\varepsilon}_{it} = \mathbf{A}_0^{-1}\cdot\mathbf{e}_{it}$ have appropriate dimensions.

In order to estimate the multivariate dynamic system (10) we employ the panel data VAR methodology (Love, 2001). This technique applies the traditional VAR approach, where all variables in the system are modelled as endogenous, within a panel data framework, where unobserved individual heterogeneity is allowed. It seems to be particularly well-suited for purposes of this study, because of the “intrinsic” endogeneity of the variables involved in the

[13] This assumption also holds when measurement errors do not exist at all and implies that $\mathrm{var}(m_{it}) = 0$ and $\tilde{m}_{it} \sim MA(1)$ otherwise.

analysis (i.e. the “same” individual firm provides all the requested information on its exporting activity). However, in using the VAR approach to panel data, we need to impose the restriction that the underlying structure is the same for each cross-sectional unit by introducing fixed effects ($\mathbf{a}_i$). Since the fixed effects are correlated with the regressors due to lags of the dependent variables, the mean differencing procedure might lead to biased coefficients estimates. To avoid this problem we use the Helmert transformation (Arellano and Bover, 1995), which removes only the forward mean, i.e. the mean of all future observations available for each sector/region-quarter. Since such a transformation preserves the orthogonality between transformed variables and lagged regressors, we use the latter as instruments and estimate the coefficients by GMM.[14] Finally, model (10) also admits region-specific time dummies ($\mathbf{d}_t$) in order to capture aggregate macro shocks that may affect all sectors in the same way. These dummies are eliminated subtracting the means of each variable calculated for each sector/region-quarter.

4.2 Estimation results: the baseline model

The starting point of the empirical analysis is the estimation of model (10). The order of autoregression is set equal to one, according to the conventional general to specific model reduction procedure.[15] Finding an exhaustive explanation for all effects of lagged variables in a vector autoregression is often both difficult (Sims, 1980) and unnecessary for the purposes of the analysis. In this work, instead, all estimated coefficients have a clear economic interpretation. Testing the structural model (9) from the reduced form (10) implies that only coefficients above the main diagonal of matrix $\mathbf{\Pi}_1$ should be not statistically significant. It translates into a conventional Granger causality test applied to each equation of the reduced form model.

Results for the full sample are shown in Table 3, which reads by column. Standard errors are reported in parentheses, while p-values are given in square brackets. Most of the autoregressive coefficients are strongly significant, underlining the existence of feedback effects among the variables of the

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[14] In this special case, the number of regressors equals the number of instruments (just-identification of the model); therefore, GMM results are numerically equivalent to those from equation-by-equation 2SLS.

[15] In all models, lagged coefficients, of order grater than one, turn to be not statistically significant at the usual significance levels.

system. Moreover, the exclusion tests indicate that the theoretical restrictions are not rejected by the data at the single equation and at the system level (lower part of the Tables), corroborating our economic priors.

Tab. 3 Baseline Panel VAR model. Estimation results: 1999*q*1-2005*q*2

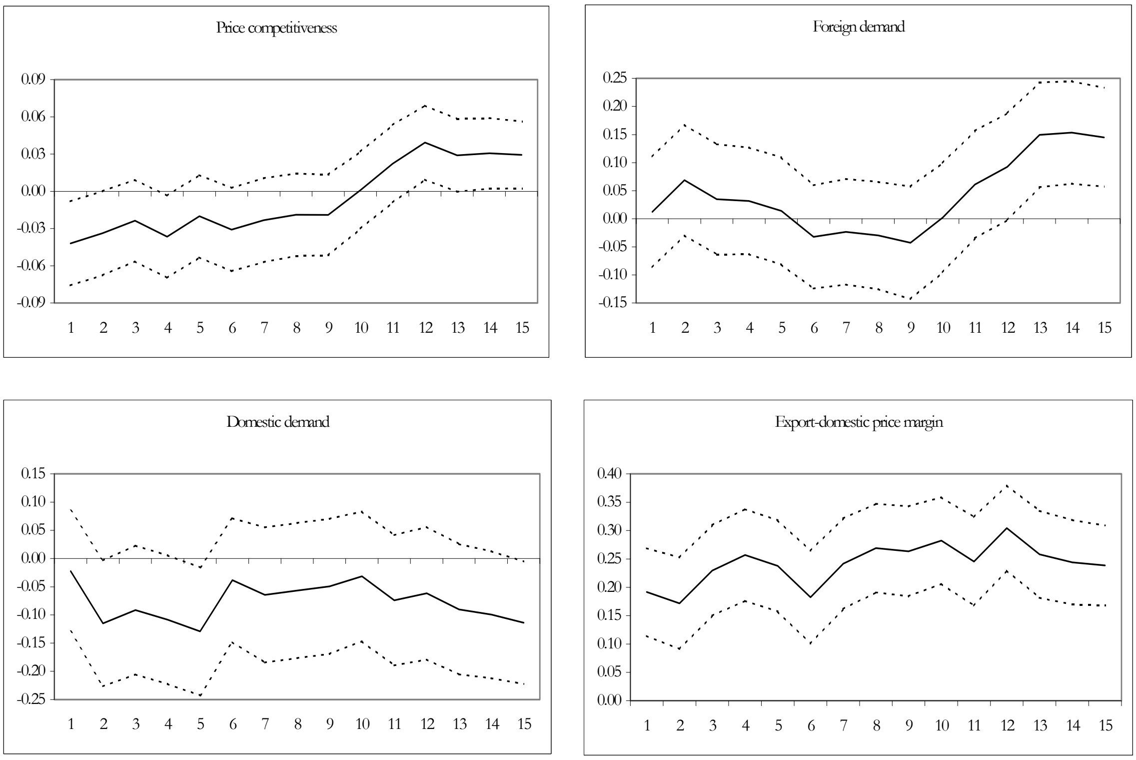
	q_t	f_t	h_t	r_t
q_{t-1}	0.3449 (0.0331)	-0.0496 (0.0127)	-0.0465 (0.0108)	-0.0037 (0.0131)
f_{t-1}	-0.1640 (0.1009)	0.5593 (0.0388)	0.1523 (0.0328)	0.0449 (0.0400)
h_{t-1}	0.0184 (0.1147)	0.0842 (0.0441)	0.4583 (0.0373)	-0.0561 (0.0454)
r_{t-1}	0.0903 (0.0804)	0.0159 (0.0309)	0.0045 (0.0262)	0.2794 (0.0318)
Number of Observations	936	936	936	936
F-test	29.08 [0.00]	119.40 [0.00]	121.92 [0.00]	19.96 [0.00]
	Exclusion Tests			
Single-equation	$f_{t-1}=h_{t-1}=r_{t-1}=0$ F(3,3728)=1.82 [0.14]	$h_{t-1}=r_{t-1}=0$ F(2,3728)=1.91 [0.15]	$r_{t-1}=0$ F(1,3728)=0.03 [0.86]	. .
System	F(6,3728)=1.55 [0.16]			

Note. The system includes price competitiveness (*q*), foreign demand level (*h*), domestic demand level (*h*) and export-domestic price margin (*r*). Standard errors (p-values) are reported in parentheses (square brackets). In single equation (system) exclusion tests, under the null hypothesis, lagged variables are deleted separately (contemporaneously) in each equation of the system.

Since the focus of the analysis is on the behaviour of the export-domestic price margins, the rest of the work is devoted to discuss the role of their determinants over time. As far as the *entire period* (1999*q*1-2005*q*2) is concerned, lagged values of *q*, *f* and *h* do not affect current values of *r*, while traces of persistence can be detected. Conversely, the statistical significance of the parameter of r_{t-1} suggests that *past* export-domestic price margins influence *current* relative prices. Predetermined variables have the expected signs, but their lack of statistical significance might suggest that market forces mainly drive the pricing behaviour of Italian exporting firms, without any room for price differentiation policies. However, that conclusion seems to stand out against the stylised facts in Section 3.

In order to investigate whether the relationship between the export-domestic price margin and its determinants is stable over time, model (10) is estimated for all sub-periods of 12 quarters. Thus, 15 sets of regressions have been run over the sample period in order to test the stability of regression coefficients.[16] Figure 3 reports the estimated coefficients (solid line) as well as ±2σ-confidence intervals (dashed lines) relative to the export-domestic price margin equation.

Fig. 3 Baseline Panel VAR model. Rolling estimation results for the export-domestic price margin equation.



Note. Estimation window consists of 12 quarters. Estimated coefficients (±2σ confidence intervals) in solid (dashed) lines.

Results from the rolling estimation indicate a strong instability of all coefficients, but the one of the export-domestic price margin. This might explain the lack of significant coefficients in Table 3. Interestingly, price competitiveness coefficients turn to be positive (negative) and statistically significant in the last estimations (first sub-samples).

The next step of the empirical analysis consists of splitting the time span into two sub-samples (1999*q*1-2001*q*4 and 2002*q*1-2005*q*2), in a way consistent with the above discussed stylised facts and with the evidence provided by the rolling estimation analysis. The estimates reported in Table 4 and 5 provide evidence of strongly different pricing behaviours.

[16] The decision to use 12-quarter sub-periods was the result of a compromise between maximizing the number of regressions and maintaining a reasonable sample size for each regression.

As far as the *first period* (1999*q*1-2001*q*4) is concerned (Table 4), while foreign and domestic demand variables turn to be statistically insignificant, the effect of price competitiveness factors results statistically significant: a 10% increase in the perceived price competition translates into a 0.4% drop for the margin between export and domestic prices. This finding seems to indicate a PTM-type behaviour of Italian exporters. In other words, firms have set prices differently in the Home and Foreign markets accordingly with the fluctuations of price competitiveness, due to both exchange rate movements and cost pressures of competitors.

Tab. 4 Baseline Panel VAR model. Estimation results: 1999*q*1-2001*q*4

	q_t	f_t	h_t	r_t
q_{t-1}	0.1990 (0.0500)	-0.0503 (0.0194)	-0.0165 (0.0171)	-0.0420 (0.0206)
f_{t-1}	-0.0780 (0.1456)	0.5293 (0.0564)	0.1613 (0.0497)	0.0122 (0.0601)
h_{t-1}	0.2670 (0.1582)	0.1116 (0.0613)	0.4387 (0.0541)	-0.0223 (0.0653)
r_{t-1}	0.1380 (0.1140)	0.0244 (0.0442)	0.0256 (0.0390)	0.1917 (0.0471)
Number of Observations	390	390	390	390
F-test	5.59 [0.00]	61.09 [0.00]	60.12 [0.00]	5.58 [0.00]
		Exclusion Tests		
Single-equation	$f_{t-1}=h_{t-1}=r_{t-1}=0$ F(3,1544)=1.80 [0.14]	$h_{t-1}=r_{t-1}=0$ F(2,1544)=1.78 [0.17]	$r_{t-1}=0$ F(1,1544)=0.43 [0.51]	. .
System		F(6,1544)=1.57 [0.15]		

Note. See Table 3.

As far as the *second period* (2002*q*1-2005*q*2) is concerned (Table 5), the estimates depict remarkable differences. Lagged demand variables have the expected signs, while the cost competitiveness indicator becomes positive, albeit statistically not significant. This finding is somewhat surprising, in that it suggests that Italian exporters did not use the price channel to offset the negative consequence of the appreciation of the euro and of the growing competitive pressure of products coming from the emerging economies (China and other Far-East countries). Moreover, foreign demand (and to a lesser extent the domestic one) seems to significantly affect export-domestic price margins, indicating a pro-cyclical pricing policy in the most recent period.

Tab. 5 Baseline Panel VAR model. Estimation results: 2002q1-2005q2

	q_t	f_t	h_t	r_t
q_{t-1}	0.2818 (0.0412)	-0.0153 (0.0155)	-0.0343 (0.0127)	0.0293 (0.0164)
f_{t-1}	0.1650 (0.1344)	0.4063 (0.0507)	0.0959 (0.0414)	0.1451 (0.0536)
h_{t-1}	-0.3159 (0.1658)	-0.0618 (0.0626)	0.2186 (0.0511)	-0.1139 (0.0662)
r_{t-1}	0.0313 (0.1073)	-0.0165 (0.0405)	-0.0477 (0.0331)	0.2384 (0.0428)
Number of Observations	468	468	468	468
F-test	13.85 [0.00]	20.78 [0.00]	16.75 [0.00]	11.94 [0.00]
		Exclusion Tests		
Single-equation	$f_{t-1}=h_{t-1}=r_{t-1}=0$ F(3,1856)=1.34 [0.26]	$h_{t-1}=r_{t-1}=0$ F(2,1856)=0.52 [0.59]	$r_{t-1}=0$ F(1,1856)=2.08 [0.15]	. .
System		F(6,1856)=1.19 [0.31]		

Note. See Table 3.

4.3 Robustness check

As a check of robustness of our results, we have applied the same econometric procedure but imposing the symmetry and proportionality condition on foreign and domestic demand variables, i.e. using a relative demand indicator (g). A closer look to the estimated demand coefficients in the baseline model reveals that lagged *h* and *f* enter in the export-domestic price margin equation with the opposite sign but similar magnitude. Over the entire EMU period, lagged foreign and domestic demand coefficients in the r_t equation are 0.0449 and -0.0561, respectively (Table 3), while in the first (second) sub-sample the estimates of f_{t-1} and h_{t-1} are 0.0122 (0.1451) and -0.0223 (-0.1139), as reported in Table 4 (Table 5). Thus, it is possible that the dynamics of export-domestic price margin actually reacts to developments in the domestic and foreign economies in a symmetric way.

Table 6 presents GMM estimates for the tri-variate panel VAR model, over the full sample. Three main findings emerge. *First*, exclusion tests indicate that the theoretical structure of reference is not rejected by the data. *Second*, the magnitude of g_{t-1} coefficient turns to be very similar to (the absolute value of) demand levels coefficients in Table 3. *Finally*, neither lagged price competitiveness nor lagged demand levels are statistically significant, with magnitudes almost identical to the ones obtained in the baseline model.

Tab. 6 Tri-variate Panel VAR model. Estimation results: 1999q1-2005q2

	q_t	g_t	r_t
q_{t-1}	0.3284 (0.0304)	-0.0074 (0.0102)	-0.0028 (0.0120)
g_{t-1}	-0.1819 (0.0879)	0.4161 (0.0295)	0.0527 (0.0348)
r_{t-1}	0.0716 (0.0756)	0.0232 (0.0254)	0.2828 (0.0299)
Number of Observations	936	936	936
F-test	40.77 [0.00]	67.97 [0.00]	31.46 [0.00]
		Exclusion Tests	
Single-equation	$g_{t-1}=r_{t-1}=0$ F(2,2799)=2.46 [0.09]	$r_{t-1}=0$ F(2,2799)=0.84 [0.36]	. .
System		F(3,2799)=1.92 [0.12]	

Note. The system includes price competitiveness (q), relative demand level (g), and export-domestic price margin (r). Standard errors (p-values) are reported in parentheses (square brackets). In single equation (system) exclusion tests, under the null hypothesis, lagged variables are deleted separately (contemporaneously) in each equation of the system.

Table 7 (Table 8) reports the results from the same estimation exercise when only the first (second) sub-sample is considered. In both regressions, exclusion *F* tests give strong support to our economic priors.

Tab. 7 Tri-variate Panel VAR model. Estimation results: 1999q1-2001q4

	q_t	g_t	r_t
q_{t-1}	0.2097 (0.0497)	-0.0315 (0.0167)	-0.0426 (0.0204)
g_{t-1}	-0.1448 (0.1402)	0.3535 (0.0471)	0.0158 (0.0577)
r_{t-1}	0.1514 (0.1139)	0.0017 (0.0383)	0.1909 (0.0469)
Number of Observations	390	390	390
F-test	6.51 [0.00]	19.74 [0.00]	7.44 [0.00]
		Exclusion Tests	
Single-equation	$g_{t-1}=r_{t-1}=0$ F(2,1161)=1.33 [0.26]	$r_{t-1}=0$ F(1,1161)=0.00 [0.96]	. .
System		F(3,1161)=0.89 [0.45]	

Note. See Table 6.

Consider the model estimated over the period 1999*q*1-2001*q*4, first. Lagged price competitiveness has a negative and statistically significant coefficient, while lagged relative demand is substantially irrelevant. Moreover, the estimated g_{t-1} coefficient turns to be very close to (the absolute value of) lagged demand levels coefficients reported in Table 4.

Tab. 8 Tri-variate Panel VAR model. Estimation results: 2002q1-2005q2

	q_t	g_t	r_t
q_{t-1}	0.2870 (0.0409)	0.0180 (0.0136)	0.0282 (0.0163)
g_{t-1}	0.2065 (0.1283)	0.3022 (0.0426)	0.1365 (0.0512)
r_{t-1}	0.0385 (0.1070)	0.0298 (0.0355)	0.2370 (0.0427)
Number of Observations	468	468	468
F-test	18.11 [0.00]	18.93 [0.00]	15.84 [0.00]
		Exclusion Tests	
Single-equation	$g_{t-1}=r_{t-1}=0$ F(2,1395)=1.46 [0.23]	$r_{t-1}=0$ F(1,1395)=0.70 [0.40]	. .
System		F(3,1395)=1.21 [0.31]	

Note. See Table 6.

Conversely, results from the most recent years show that the g_{t-1} coefficient is statistically significant, with a magnitude (0.1365) roughly nine times the one estimated over the period 1999*q*1-2001*q*4 (0.0158). Furthermore, the price competitiveness indicator enters with the positive sign, albeit it is only borderline statistically significant (10% level).

5 IMPULSE RESPONSE ANALYSIS

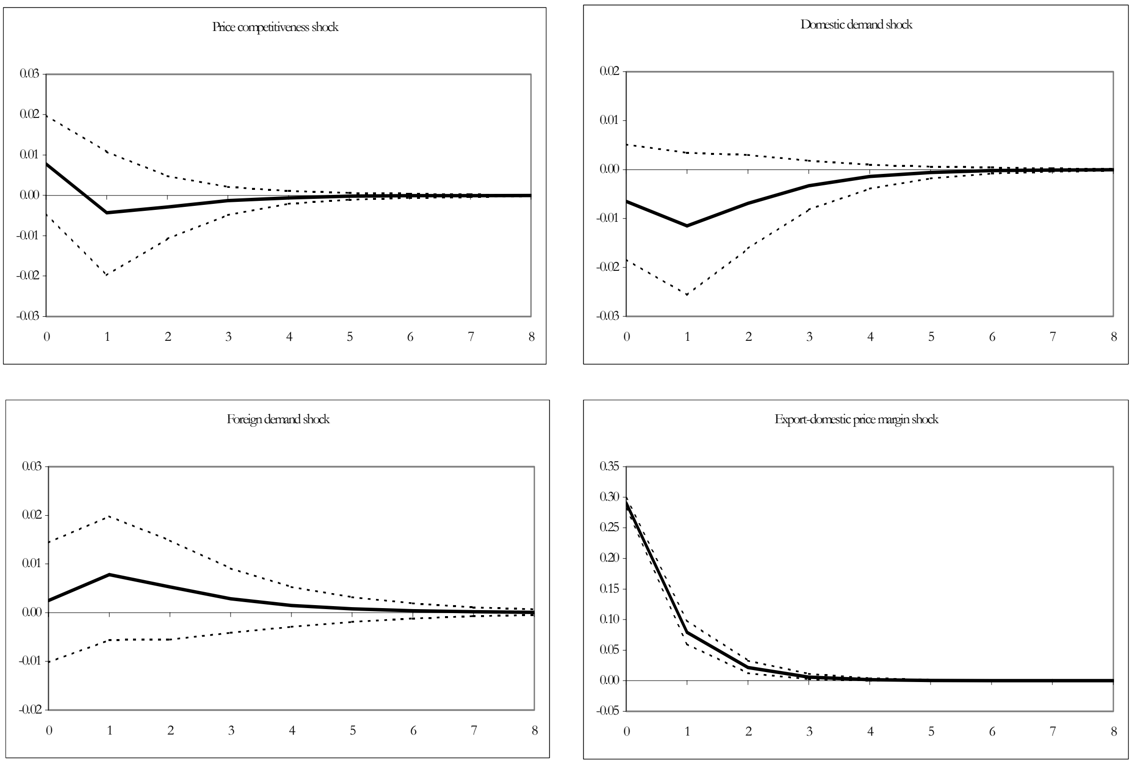
Estimated models can be used to assess the relative contribution of mutually orthogonal (structural) factors in explaining the dynamics of export-domestic price margins to support previous interpretations. In the following, structural residuals are extracted from reduced-form disturbances through an identification scheme based on the recursive structure discussed in Section 4.1. Dynamic simulations are based on the analysis of impulse response functions

(IRFs), which allow revealing the dynamic effects of one-time structural shock on the export-domestic price margin. The simulation horizon is set equal to eight quarters.

5.1. Evidence from the baseline model

Figure (4) illustrates the response of relative prices (solid line) to a positive shock to price competitiveness, foreign demand, domestic demand, and export-domestic price margin from the model estimated over the full sample period. Confidence bounds at the 95% significance level (dashed lines) are also reported.

Fig. 4 Baseline Panel VAR model. Impulse response analysis: 1999q1-2005q2

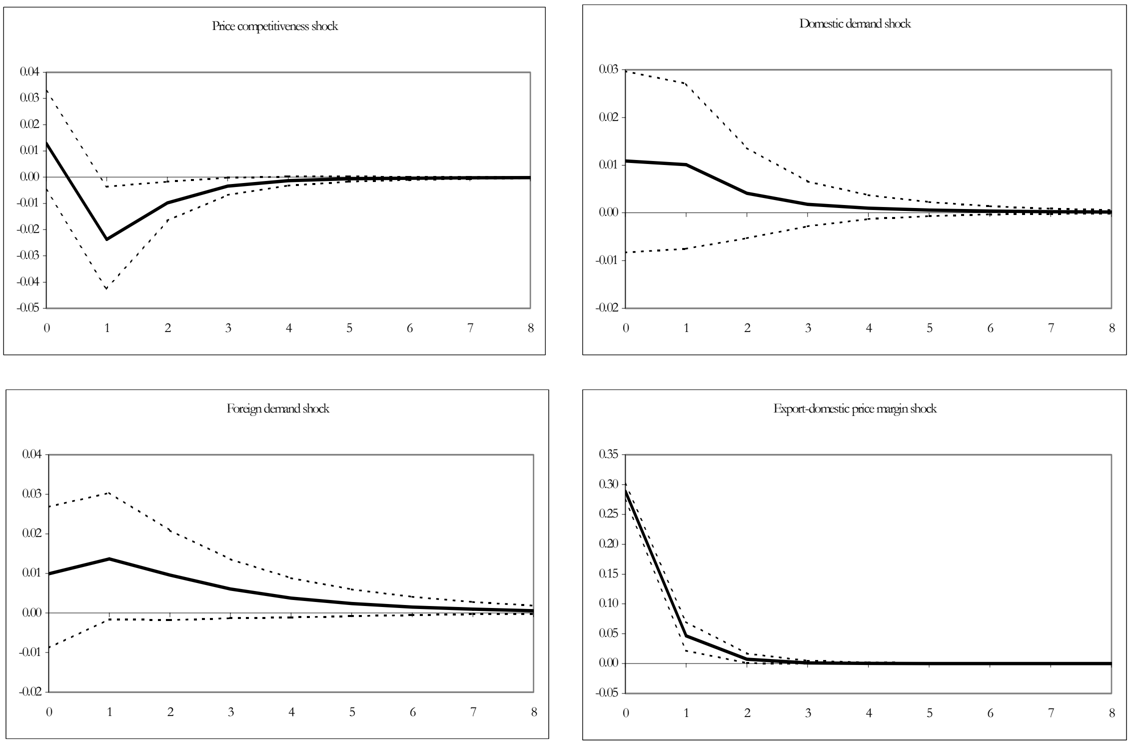


Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

The contraction (increase) of the export-domestic price margin to domestic (foreign) demand and to price competitiveness shocks is consistent with the expected economic relationships. Since confidence bounds include the baseline path (the horizontal axis), deviations from the pre-shock level cannot be judged as statistically relevant at the chosen significance level. Conversely, an export-domestic price margin shock produces some effects only in the very short-run, with a half-life of the response equal to one quarter.[17]

Graphs in Figure (5) refer to the impulse response functions calculated for the first sub-period (1999*q*1-2001*q*4).

Fig. 5 Baseline Panel VAR model. Impulse response analysis: 1999q1-2001q4



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

Consider the effects of a price competitiveness shock, first. The export-domestic price margin reacts negatively one quarter after the shock; then, the response decreases over time to fully disappear within the first year of simulation. Foreign and domestic demand shocks produce negligible effects, while the response to an export-domestic price margin shock is fully absorbed within the third quarter of the simulation horizon.

Finally, Figure (6) presents the impulse response functions relative to the second sub-period.

The simulation exercise produces results rather different from those in Figure 5. While the export-domestic price margin remains substantially inelastic to changes in the domestic demand, a foreign demand shock shifts upward the export-domestic price margin in the first year of simulation. The response is

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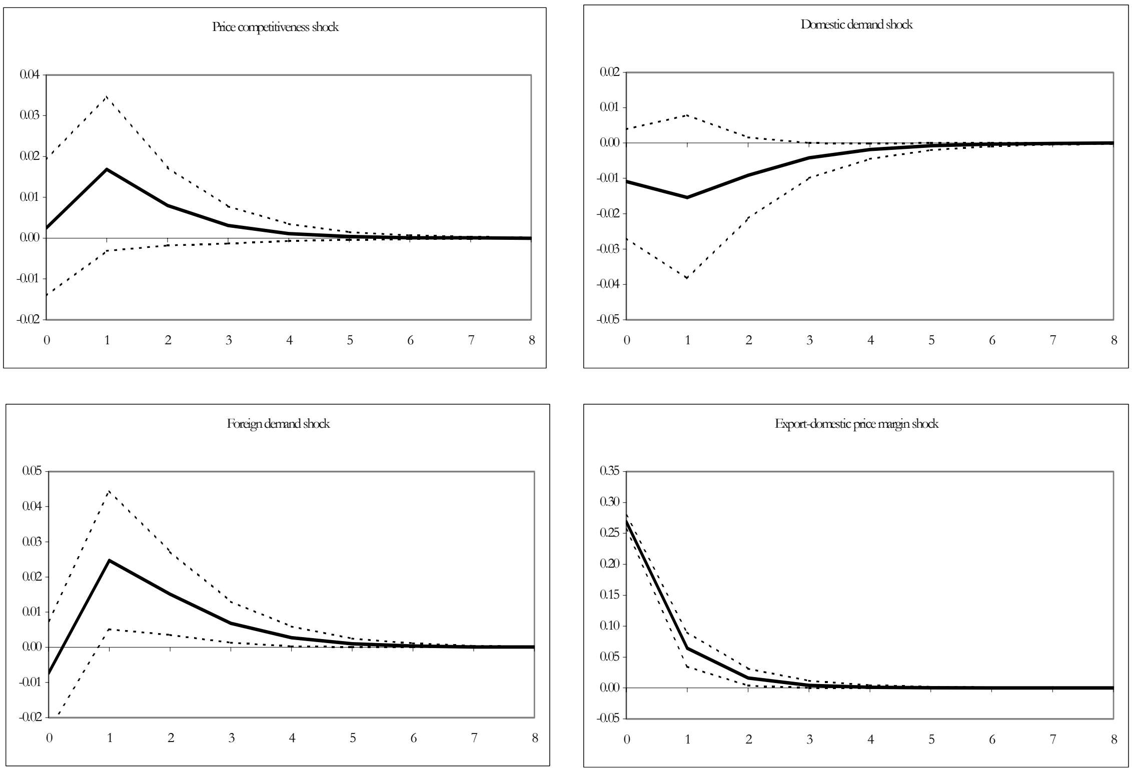
17 Half-life is defined as the number of quarters which have to pass before the deviation from the steady-state falls to half the size of the initial shock.

statistically significant and a certain degree of sluggishness in the deviation relative from the pre-shock level can be detected. More interestingly, the effects of a cost competitiveness shock are now positive (but not statistically significant). This would confirm that, during the period 2002*q*1-2005*q*2, Italian exporters' pricing policy was mainly influenced by cyclical conditions of foreign demand rather than by exchange rate and price competitiveness factors.

Fig. 6 Baseline Panel VAR model. Impulse response analysis: 2002q1-2005q2



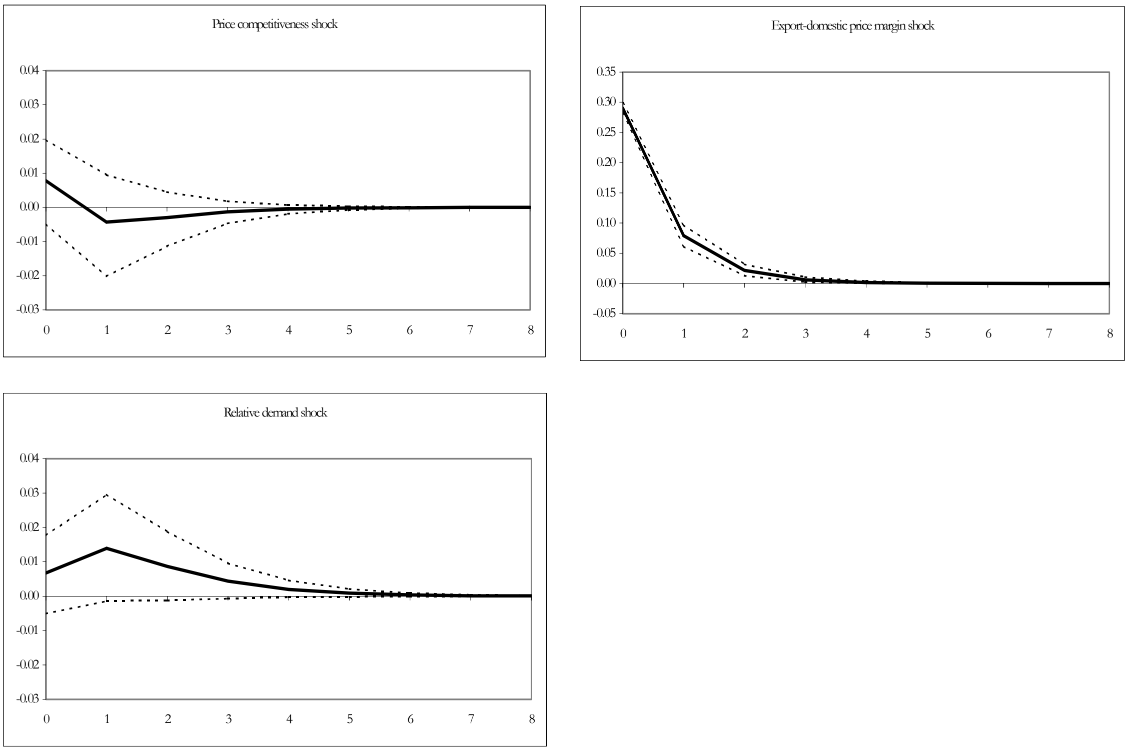
Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

5.2 Evidence from the trivariate model

Figure 7 (Figure 8 and 9) shows the response of export-domestic price margin from the tri-variate system estimated over the full sample (first and second sub-period, respectively). The overall picture emerging from the graphs is widely consistent with the conclusions discussed in Section 4.3. When the entire sample is considered, the export-domestic price margin reacts only after a shock to itself, while neither relative demand nor price competitiveness shock produces significant reactions. Dissecting the analysis in two sub-periods allows shedding light onto pricing behaviour of Italian exporting firms. Focusing on the first sub-sample, the indications from the IRFs relative to the baseline model (Figure 5) are fully supported by the simulations from its tri-variate counterpart

(Figure 8). Conversely, the evidence relative to the most recent years indicates that the response of export-domestic price margin is positive and statistically significant not only to a relative demand shock but also to a price competitiveness shock.[18]

Fig. 7 Tri-variate Panel VAR model. Impulse response analysis: 1999q1-2005q2



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

IRFs analysis from the baseline and the tri-variate model allows us to conclude that unanticipated changes in cost competitiveness and *relative* demand levels appear to exert non-negligible effects on export-domestic price margins. While a typical PTM behaviour emerges over the period 1999*q*1-2001*q*4, the influence of cyclical demand conditions seem to prevail during the most recent years. A possible explanation of this quite odd behaviour can be related to an upgrading of the price/quality mix of goods sold abroad following fiercer price competition from low-cost producers and an appreciated exchange rate (the “qualitative barrier” discussed in de Nardis and Pensa, 2004). Such pro-cyclical price policies on foreign markets could have acted as a buffer to compensate the fall of profits in the domestic market, characterized by an extremely weak demand in the last years.

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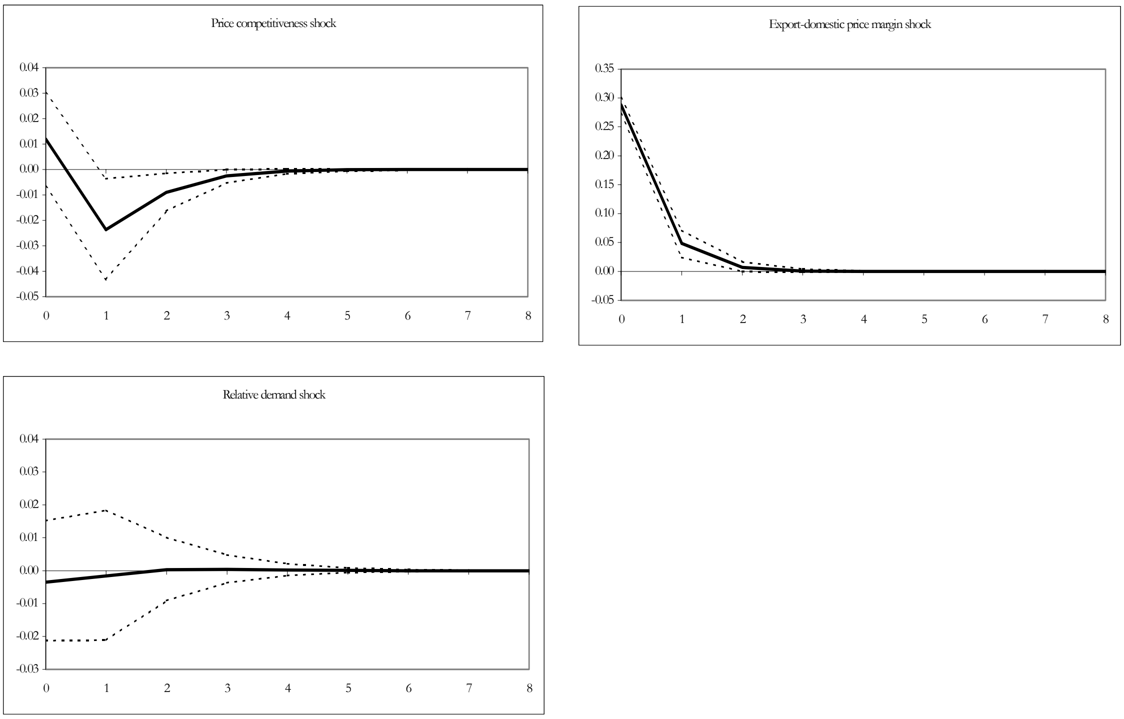
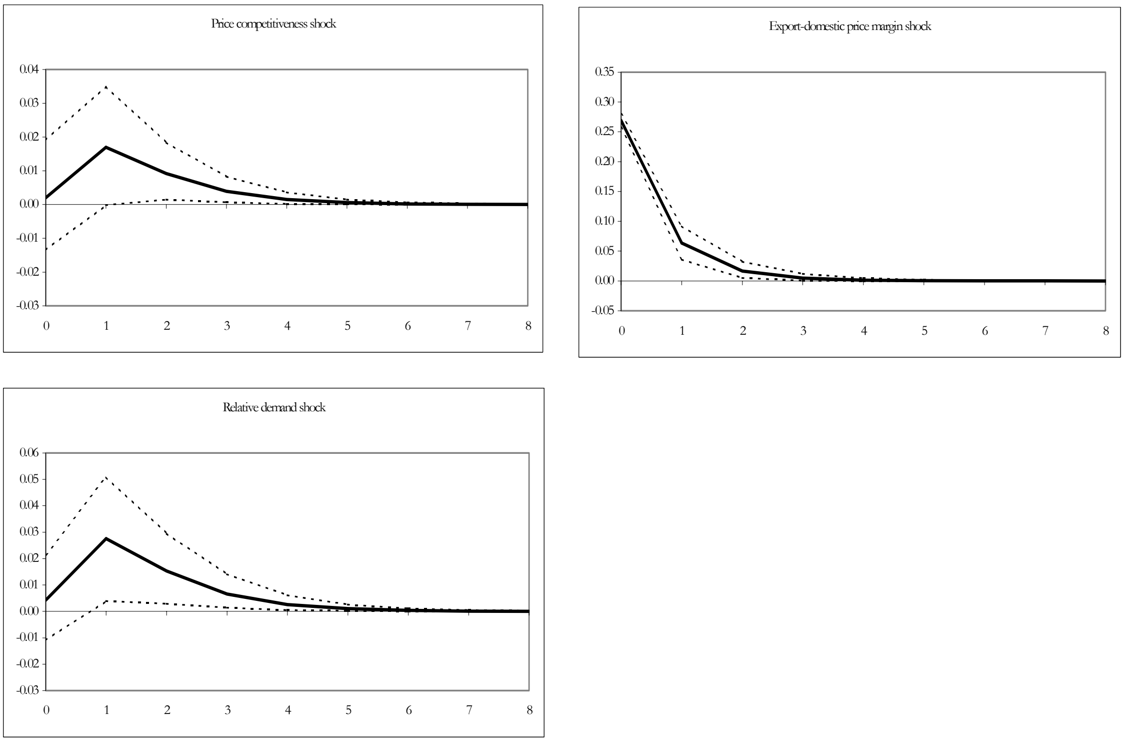
[18] Lagged reactions to price competitiveness shocks (Figures 5, 6, 8 and 9) might be considered as consistent with models based on sticky prices adjustment due to menu-cost-driven pricing behaviour or other nominal rigidities (see, among others, Ghosh and Wolf, 1994).

Fig. 8 Tri-variate Panel VAR model. Impulse response analysis: 1999q1-2001q4



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

Fig. 9 Tri-variate Panel VAR model. Impulse response analysis: 2002q1-2005q2



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

5.3 Evidence from a more disaggregate analysis

We further discuss the determinants of pricing policies in order to test a possible heterogeneity across sectors. In doing that, we distinguish between traditional sectors (i.e. those usually referred as “Made-in-Italy”, which include sub-sections DB, DC, DI e DN indicating respectively textile products, leather products, non metallic mineral products and furniture) and other sectors (i.e. the remaining sectors included in Table 1). Accordingly, our datasets consist of: *1)* a panel of 12 cross-sectional observations (4 sectors and 3 regions) and 26 time points for the “traditional sectors” model; *2)* a panel of 27 cross-sectional observations (9 sectors and 3 regions) and 26 time points for the “other sectors” model. We compare IRFs from the two models and calculate their differences (“traditional sectors” minus “other sectors”) with respect to the sub-samples 1999*q*1-2001*q*4 and 2002*q*1-2005*q*2.[19]

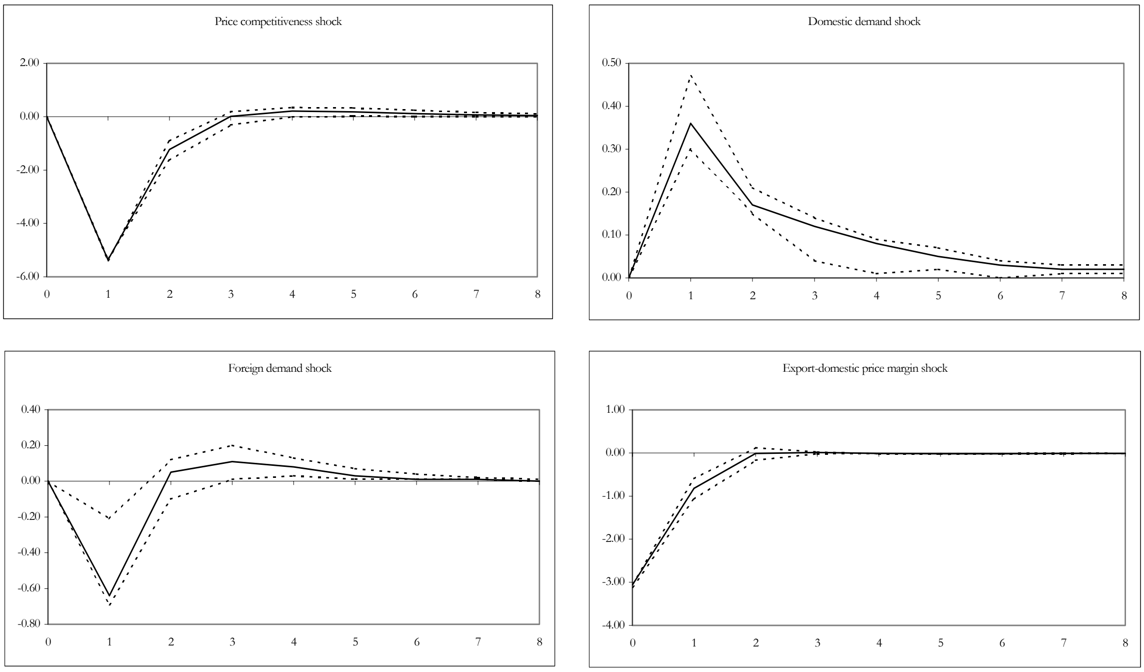
Results from the two models[20] indicate that a deterioration of price competitiveness translates into a decrease (increase) of the export-domestic price margin in the first (second) sub-sample, while the time profile of the responses relative to other shocks are consistent with our economic priors. Figure 10 (11) presents the difference of the responses of the export-domestic price margin to shocks to price competitiveness, foreign demand, domestic demand, and export-domestic price margin from the models estimated over the first (second) sub-sample.

We observe a strong heterogeneity not only between the two sub-periods, as discussed previously, but also across sectors. More in details, the increase in export-domestic price margins consequent, in the second sub-period, to a rise in cost competitiveness obstacles turns to be higher in the “traditional sectors” model. Conversely, the export-domestic price margins reactions in the “other sectors” model appear to be more elastic to unanticipated changes in demand levels, albeit the results relative to a domestic demand shock are not significant in the second sub-sample, suggesting a homogeneous reaction in both of models. As suggested by de Nardis and Pensa (2004), possible explanations of this rely on displacement of uncompetitive producers and survival of exporting firms able to make price in international markets. As a result, surviving exporters in these industries were those “shielded” from foreign competition thanks to quality differentiation and who were able to adopt pricing policies in international markets partly independent of competitors’ behaviour.

[19] Given that the two samples are independent, IRFs of the differences are equivalent to the difference of IRFs. The same argument holds for the confidence bounds.

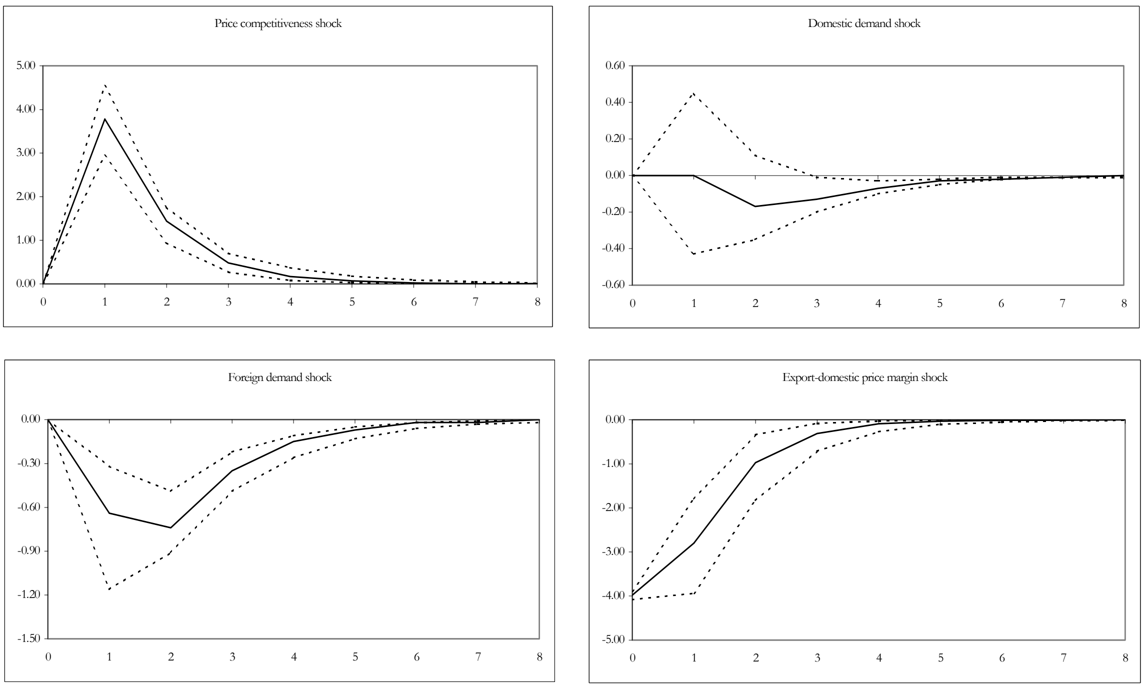
[20] Not reported for sake of brevity.

Fig. 10 Sector Panel VAR models. Differences in impulse-responses ("Traditional sectors" minus "Other sectors"): 1999q1-2001q4 Percentage values



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

Fig. 11 Sector Panel VAR models. Differences in impulse-responses ("Traditional sectors" minus "Other sectors"): 2002q1-2005q2 Percentage values



Note. 95% confidence bounds (dashed lines) are generated by Monte-Carlo with 1000 replications.

5.4 An overall discussion: from a firm-level to a macroeconomic perspective

A discussion focused only on a firm-based level may tell part of the whole story. Looking at the macroeconomic consequences of the observed pricing policy in our panel of firms allows interpreting the results in (at least) two additional ways. *First*, PTM policies are often evoked as a primary source of the empirical rejection of the LOP (Engel, 1993). The empirical evidence presented in Section 4 and 5 seems to contrast with this view. Short-lived deviations from the baseline path emerging from the IRFs exercise suggest that PTM policies may be pursued only in the very short-run. Moreover, there are no statistically significant responses of the export-domestic price margin to the various shocks when a longer simulation horizon is taken into account. Finally, the extent of the deviations from the baseline path in the IRFs is indeed well below the level predicted by theoretical models based on translog preferences, such as the one proposed by Bergin and Feenstra (2001). As shown in Haskel and Wolf (2001), differences in local distribution costs, local taxes, and tariffs as well as in other additional factors may allow firms to vary mark-ups over time, although the convergence between prices may be restored in the long-run. In this sense, the results from our study conforms recent findings on threshold mean reversion for sector goods (see Obstfeld and Taylor, 1997; Sarno and Taylor, 2002). *Second*, a reduced incompleteness of ERPT associated to a limited PTM should have no dramatic consequences on the classical textbook reaction of the current account to exchange rate modifications based on the Marshall-Lerner requirements on export and import price elasticities. Under this point of view, the results of the present study are consistent with the evidence of improvements of the trade balance to real exchange rate depreciations as discussed in Boyd, Caporale and Smith (2001) for the G7 economies.

6 CONCLUSIONS

We revisited the model proposed by Marston (1990) to test PTM of Italian exporting firms over the period 1999*q*1-2005*q*2. We looked at the average behaviour of a sample of Italian exporting industrial firms, during a very critical period (the EMU years, since 1999) when a significant reduction of the aggregate Italian market share took place in volume terms.

The use of survey data made it possible to use as relevant variable in the analysis an indicator of the export-domestic price margin for the "same" good produced by the "same" firm, permitting to capture effective price differences from two different transactions made by the same production unit. This is a significant step forward with respect to an analysis based on export unit values, which are affected by unavoidable composition effects and are not consistent with the domestic price indicator adopted as comparison (usually, producer prices). The empirical investigation used panel data VAR models, which treat all variables in the system as endogenous within a framework allowing for unobserved individual heterogeneity. Impulse-response analysis helped assessing the reaction of export-domestic price margins to unanticipated changes in cost competitiveness and demand levels.

These factors appeared to exert non-negligible effects. Specifically, for the period 1999*q*1-2001*q*2 a typical PTM behaviour emerged, while during the period 2002*q*1-2005*q*2 the influence of cyclical demand conditions prevailed, inducing exporting Italian firms to increase their export-domestic price margins to take advantage of a more favourable foreign demand in face of a strong deterioration of their cost competitiveness. However, IRFs exercises showed low persistent reactions of variables once shocked, suggesting short-lived PTM policies. Moreover, taking into account sector heterogeneity, we found indications that the anomalous price setting in the second sub-period was mainly due to traditional ("Made-in-Italy") sectors, i.e. those most hit by external competitive pressures.

From a macroeconomic point of view, consequences of a limited PTM (not only in timing but also in intensity) are at least twofold. First, PTM policies should not be relevant in explaining long-run deviations from the LOP. Second, no dramatic effects on the classical textbook reaction of the current account to exchange rate modifications should be observed. Both results are consistent with the most recent empirical literature.

APPENDIX

The elasticity of the export-domestic price margin is given by

$$\text{(A1)}\quad \varepsilon_{(P_{it}^F/P_{it}^H)E_t} = \frac{\partial R_{it}}{\partial E_t}\cdot\frac{E_t}{R_{it}} = \varepsilon_{P_{it}^F E_t} - \varepsilon_{P_{it}^H E_t}$$

Indicating by M_{it}^F the variable mark-up over marginal cost in the Foreign market, explicit formulations of the elasticities on the R.H.S. of (A1) are the following

$$\varepsilon_{P_{it}^F E_t} = \frac{-\varepsilon_{M_{it}^F(P_{it}^F/E_t)} - \varepsilon_{MC_{it}X_{it}^F}\cdot\varepsilon_{X_{it}^F(P_{it}^F/E_t)}\cdot\varepsilon_{P_{it}^F MC_{it}}}{1-\varepsilon_{M_{it}^F(P_{it}^F/E_t)} - \varepsilon_{MC_{it}X_{it}^F}\cdot\varepsilon_{X_{it}^F(P_{it}^F/E_t)}\cdot\varepsilon_{P_{it}^F MC_{it}}}$$

$$\varepsilon_{P_{it}^H E_t} = \varepsilon_{MC_{it}X_{it}^F}\cdot\varepsilon_{X_{it}^F(P_{it}^F/E_t)}\cdot(\varepsilon_{P_{it}^F E_t}-1)\cdot\varepsilon_{P_{it}^H MC_{it}}$$

where $\varepsilon_{M_{it}^F(P_{it}^F/E_t)}$ is the elasticity of the mark-up practiced in the Foreign market with respect to the (foreign currency) price made in that market (P_{it}^F/E_t).

This elasticity is negative, provided the demand schedule is less convex than a constant elasticity curve (such that the mark-up reduces when the price rises and the quantity falls). As for the other terms in the expressions, $\varepsilon_{MC_{it}X_{it}^F}$ is the (positive) elasticity of marginal costs with respect to foreign sales, $\varepsilon_{X_{it}^F(P_{it}^F/E_t)}$ is the (negative) elasticity of foreign sales with respect to foreign currency prices practiced in the Foreign market, and $\varepsilon_{P_{it}^F MC_{it}}$ and $\varepsilon_{P_{it}^H MC_{it}}$ are the (positive) elasticities of (domestic currency) prices practiced in Foreign and Home markets with respect to the (common) marginal cost. Given these signs, both $\varepsilon_{P_{it}^F E_t}$ and $\varepsilon_{P_{it}^H E_t}$ are positive (and less than 1), indicating positive reactions of the (domestic currency) prices in the Foreign and Home markets to exchange rate changes.

As to the sign of the elasticity of the export-domestic price margin to exchange rate changes, expression (A1) can be written as

$$\varepsilon_{(P_{it}^F/P_{it}^H)E_t} = \frac{-\varepsilon_{M_{it}^F(P_{it}^F/E_t)} - \varepsilon_{MC_{it}X_{it}^F}\varepsilon_{X_{it}^F P_{it}^F/E_t}(\varepsilon_{P_{it}^F MC_{it}} - \varepsilon_{P_{it}^H MC_{it}})}{(1-\varepsilon_{M_{it}^F P_{it}^F/E_t} - \varepsilon_{MC_{it}X_{it}^F}\varepsilon_{X_{it}^F P_{it}^F/E_t}\varepsilon_{P_{it}^F MC_{it}})(1-\varepsilon_{MC_{it}X_{it}^F}\varepsilon_{X_{it}^F P_{it}^F/E_t}\varepsilon_{P_{it}^H MC_{it}})}$$

which is positive and less than 1 when

$$-\varepsilon_{M_{it}^F(P_{it}^F/E_t)} > \varepsilon_{MC_{it}X_{it}^F}\varepsilon_{X_{it}^F P_{it}^F/E_t}(\varepsilon_{P_{it}^F MC_{it}} - \varepsilon_{P_{it}^H MC_{it}})$$

This is certainly the case when the export and the domestic prices, expressed in domestic currency, respond in the same way to marginal cost changes. In these circumstances, the margin between foreign and domestic prices rises (falls) by a fraction of the increase (reduction) of the exchange rate.

It is clear from (A1) and the following definitions for $\varepsilon_{P_{it}^F E_t}$ and $\varepsilon_{P_{it}^H E_t}$ that when marginal cost is independent of output (so that there are no motives for the price practiced in the Home market to vary after an exchange rate modification), the elasticity of the export-domestic price margin with respect to exchange rate becomes

$$\varepsilon_{(P_{it}^F / P_{it}^H) E_t} = \frac{-\varepsilon_{M_{it}^F (P_{it}^F / E_t)}}{1 - \varepsilon_{M_{it}^F (P_{it}^F / E_t)}}$$

which is positive and less than 1. Moreover when mark ups are invariant with respect to prices (and, hence, exchange rate), the elasticity of the export-domestic price margin is null. In this case, whether ERPT is complete or not depends on the behaviour of marginal costs.

If the marginal cost rises with output then P_{it}^F and P_{it}^H vary by the same amount following an exchange rate change, since

$$\varepsilon_{P_{it}^F E_t} = \varepsilon_{P_{it}^H E_t} = \frac{-\varepsilon_{MC_{it} X_{it}^F} \cdot \varepsilon_{X_{it}^F (P_{it}^F / E_t)}}{1 - \varepsilon_{MC_{it} X_{it}^F} \cdot \varepsilon_{X_{it}^F (P_{it}^F / E_t)}}$$

which is positive and less than 1; so that, with constant mark-ups and marginal cost increasing in output ERPT is incomplete.

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