ELSEVIER

Journal of International Economics 65 (2005) 151–165

Journal of INTERNATIONAL ECONOMICS

www.elsevier.com/locate/econbase

Antidumping protection and markups of domestic firms

Jozef Konings, Hylke Vandenbussche*

Faculty of Economics and Licos, Catholic University of Leuven, Naamsestraat 69, 3000 Leuven, Belgium
CEPR, London, UK

Received 24 September 2002; received in revised form 3 February 2004; accepted 26 February 2004

Abstract

This paper tests whether Antidumping (AD) protection affects the market power of import-competing domestic firms. We use panel data of about 4000 EU producers that were involved in AD cases to estimate markups before and after the filing of a case. Our findings indicate that AD protection has positive and significant effects on domestic markups, except in cases where import diversion after protection is strong, like in 'seamless steel tubes'. Our results control for potential endogeneity of AD filings. A randomly drawn control group of firms not subject to AD policy did not have rising markups during the same period.

Keywords: Markups; Antidumping protection; European producers; Firm data

JEL classification: F13; L13; L41

1. Introduction

Among trade economists, there is a growing consensus that in many cases, Antidumping (AD) policy is an industrial policy tool in disguise. Instead of keeping 'unfair imports' out, it is often aimed at fostering the interests of domestic producers (Blonigen and Prusa, 2003). However, despite the industrial policy nature of AD measures, surprisingly little empirical work has measured the effects of AD policy on domestic producers.[1] In this

* Corresponding author. Tel.: +32-16-32-69-20; fax: +32-16-32-67-32.
E-mail address: hylke.vandenbussche@econ.kuleuven.ac.be (H. Vandenbussche).

[1] A small number of papers have used US stock market data and identified excess returns on firms protected by trade policy (e.g., Lenway et al., 1990).

0022-1996/$ - see front matter © 2004 Elsevier B.V. All rights reserved.
doi:10.1016/j.jinteco.2004.02.001

paper, we look at how domestic producers' markups, defined as price over marginal cost, are affected by AD protection. This seems a natural focus, given that many existing theoretical models of AD have analyzed domestic profitability and price setting behavior.

In theory, an AD duty is very similar to an import tariff.[2] The static effects on an imperfectly competitive industry of a tariff/duty on foreign imports point in the direction of a rise in domestic prices irrespective of the type of competition assumed (Helpman and Krugman, 1989). Several papers have also pointed out that AD protection can result in collusive outcomes with higher prices in the domestic market for both domestic and foreign firms involved in the AD case.[3] A number of dynamic models have also been developed, taking into account that firms involved in AD cases may have incentives to behave strategically to influence AD outcomes.[4] While these models offer some guidance as to what we may expect a priori about the direction of markups, it is not our intention to test any of these models formally. Rather, our primary focus will be to test for a structural break in domestic markups as a result of AD protection.

In particular, we will study the 1996 European AD cases and the domestic producers affected by them. Firm level data for the period 1992–2000 will allow us to estimate markups before and after AD protection for this group of domestic producers.[5]

Given that earlier empirical work has shown that trade liberalization disciplines markups (Levinsohn, 1993; Harrison, 1994; Krishna and Mitra, 1998), we are inclined to expect, a priori, that trade protection will raise markups. However, there are a number of reasons why the effect of AD protection on markups may not be unambiguously positive. Import diversion from dumping countries to non-dumping countries (Prusa, 1997), domestic entry and/or inward FDI[6] are just a few reasons why the increase of domestic markups due to AD protection could be dampened.

Section 2 explains the methodology we apply and discusses the company data that we use. In Section 3, we discuss our findings both for the pooled data across AD cases as well as on a case-by-case basis. In Section 4, we perform a number of extra robustness checks, and Section 5 is a concluding one.

2. Empirical methodology and data

2.1. Methodology

There are many alternative ways to estimate markups.[7] Any choice between them is likely to involve trade offs. Our methodology for estimating markups is based on Roeger

[2] In the EU, AD measures can also take the form of a price undertaking (PU), which is voluntary price increase by the importers and is also believed to raise domestic markups (Belderbos et al., 2004).

[3] For example, Veugelers and Vandenbussche (1999) and Zanardi (2004).

[4] For example, Fischer (1992), Reitzes (1993), Prusa (1994) and Pauwels et al. (2001).

[5] Under EU AD law, protection is in principle limited to 5 years (Sunset Clause).

[6] We do not observe entry and exit of firms. An 'entrant' in the data may just be a firm that meets the inclusion criteria in our data.

[7] For an overview on how to estimate markups with firm data, see Tybout (2003).

(1995), which is well suited for the firm level data we use. We give a brief summary of the approach below; for detailed derivations we refer the reader to Roeger (1995).

Let the output of the ith firm in year t (Q_{it}) be determined by the production function

$$Q_{it} = \Theta_{it} F(N_{it}, K_{it}, M_{it}), \tag{1}$$

where $F(N_{it},K_{it},M_{it})$ is linear homogeneous in labor (N_{it}), capital (K_{it}) and materials (M_{it}), and Θ_{it} is a firm- and period-specific productivity shock.[8] Under imperfect competition, the primal Solow residual, SR_{it}, can be decomposed into an imperfect competition term (a) and a productivity term (b) as follows (Hall, 1988):

$$\mathrm{SR}_{it} = \hat{Q}_{it} - \alpha_{Nit}\hat{N}_{it} - \alpha_{Mit}\hat{M}_{it} - (1-\alpha_{Nit}-\alpha_{Mit})\hat{K}_{it} = \underbrace{\beta_{it}(\hat{Q}_{it}-\hat{K}_{it})}_{\text{(a)}} + \underbrace{(1-\beta_{it})\hat{\Theta}_{it},}_{\text{(b)}} \tag{2}$$

where α_{Jit}=($P_{Jit}J_{it}/P_{it}Q_{it}$), J=N,M, denotes factor shares in sales, and cavats denote growth rates. Market power is captured by the Lerner index, β_{it}=($P_{it}-c_{it}/P_{it}$)=1 − (1/μ_{it}), where c_{it} stands for the marginal cost of firm i at time t, P_{it} is the product output price, and μ_{it}=(P/c)$_{\mathrm{it}}$ is the price–cost markup. In Eq. (2), the output and input factors and the factor shares can be observed from the data, but the Lerner index and the productivity shocks cannot.

The productivity shock, Θ_{it}, can be thought of as consisting of two parts:

$$\Theta_{it} = \lambda_{it} + \psi_{it}. \tag{3}$$

Firms cannot anticipate the realization on the second component, ψ_{it}, so it can be considered as white noise. However, they can anticipate the first component, λ_{it}, and it may therefore be correlated with inputs. Since both components are unobservable to the econometrician, the application of ordinary least squares to Eq. (2) is likely to result in simultaneity bias.

To deal with this problem, Roeger (1995) derives the dual or price-based Solow residual, with R_{it} referring to the rental price of capital,

$$\begin{aligned}\mathrm{DSR}_{it} &= \alpha_{Nit}\hat{P}_{Nit} + \alpha_{Mit}\hat{P}_{Mit} + (1-\alpha_{Nit}-\alpha_{Mit})\hat{R}_{it} - \hat{P}_{it}\\ &= -\beta_{it}(\hat{P}_{it}-\hat{R}_{it}) + (1-\beta_{it})\hat{\Theta}_{it}\end{aligned} \tag{4}$$

By subtracting Eq. (4) from (2), he obtains the *net* Solow residual or

$$\begin{aligned}\mathrm{SR}_{it} - \mathrm{DSR}_{it} &= (\hat{Q}_{it}+\hat{P}_{it}) - \alpha_{Nit}(\hat{N}_{it}+\hat{P}_{Nit}) - \alpha_{Mit}(\hat{M}_{it}+\hat{P}_{Mit})\\ &\quad - (1-\alpha_{Nit}-\alpha_{Mit})(\hat{K}_{it}+\hat{R}_{it}) = \beta_{it}[(\hat{Q}_{it}+\hat{P}_{it}) - (\hat{K}_{it}+\hat{R}_{it})]\end{aligned} \tag{5}$$

[8] The constant returns to scale assumption could bias downwardly the estimated changes in the markup. For a discussion see Konings and Vandenbussche (2004).

Note that the term causing the endogeneity problem, $(1-\beta_{it})\hat{\Theta}_{it}$, does not appear in Eq. (5).[9] This implies that Eq. (5) can be estimated consistently without having to use instrumental variables (IV), which are often difficult to find in micro data.[10]

We can also rewrite Eq. (5) to obtain a direct measure of the price–cost markup, $(P/c)it=\mu_{it}$,

$$\begin{aligned}\left(\hat{Q}_{it}+\hat{P}_{it}\right)-\left(\hat{K}_{it}+\hat{R}_{it}\right)=\mu_{it}\{&\alpha_{Nit}\left[\left(\hat{N}_{it}+\hat{P}_{Nit}\right)-\left(\hat{K}_{it}+\hat{R}_{it}\right)\right]\\&+\alpha_{Mit}\left[\left(\hat{M}_{it}+\hat{P}_{Mit}\right)-\left(\hat{K}_{it}+\hat{R}_{it}\right)\right]\}\end{aligned}\qquad(6)$$

Despite its apparent complexity, Eq. (6) can easily be estimated with firm level data in order to arrive at an estimate for the markup coefficient μ. The single bracketed terms in Eq. (6) all refer to growth rates of nominal values of output and input factors. Therefore, the data requirements are limited to sales $(P_{it}Q_{it})$, the wage bill of workers $(P_{Nit}N_{it})$, the nominal value of the material costs $(P_{Mit}M_{it})$ and the nominal value of capital $(R_{it}K_{it})$. The advantage of using nominal variables is that they do not need to be deflated.

For capital, we use the book value of the fixed tangible assets. For the rental price of capital (R_{it}), we follow Hall and Jorgenson (1967) and Hsieh (2002), or $R_{it}=P_I(r_t+\delta_{it})$, where P_I stands for the index of investment goods prices, measured at the country level,[11] r_t stands for the real interest rate for each period t for the country the firm belongs to and δ stands for the firm level depreciation rate on fixed tangible assets which we assume to be 10%.[12]

To simplify notation, we denote the left-hand side in Eq. (6) by ΔY_{it}, which can be interpreted as the growth rate in sales per value of capital in firm i. The terms in between brackets on the right hand side of Eq. (6) are denoted by ΔX_{it} and can be interpreted as a composite variable that represents the growth rates in the various input factors weighted by their respective share in sales. For empirical tractability, we assume, as is done in most applications of this type, that markups are the same for all firms within the same sector.

Given the short-time dimension of the data, we do not have sufficient degrees of freedom to estimate a markup for each firm separately. The full model is given in Eq. (7) where we interact the composite variable ΔX_{it} with an AD dummy equal to 1 for the years during which AD protection applies (from 1997 onwards) and 0 otherwise. In addition, we also interact ΔX_{it} with yearly GDP growth per EU country k for the country where firm i is located, to control for changes in markups due to business cycle fluctuations, demand and time effects (e.g., Rotemberg and Woodford, 1991; Roeger, 1995).

[9] This approach can be compared to methods for estimating total factor productivity consistently, as in Olley and Pakes (1996) and Pavcnik (2002).

[10] Arellano and Bond (1991) use lagged values of the endogenous explanatory variables as instruments. This approach requires a large cross-section dimension of the data to avoid finite sample bias, making its use less feasible on a case-by-case basis. But in Section 4, we report the results of an instrumental variable approach on our pooled data as a robustness check and to control for potential measurement error in the variables that can result in a non-zero error term.

[11] This variable comes from the AMECO-database from the European Commission.

[12] We experimented with depreciation rates of 15–20%, but results remained the same.

Table 1
European antidumping cases initiated in 1996

Product	Trade-weighted duty	Import share dumpers[a] (1996)	Number of EU firms in final sample[b]	Number of initiating firms
Cotton fabrics	19.6%	53%	136	5
Synthetic fibre ropes	Termination	14%	188	1
Luggage and travel goods	Termination	79%	1510	–
Leather handbags	39%	46%	1120	2
Farmed Atlantic salmon	4% + PU[c] for some firms	88%	417	14
Seamless steel pipes and tubes	27.9%	77%	114	8
Polyester fibre yarns	15%	17%	82	7
Bed linen	16%	58%	21	13
Video tapes	Termination	34%	21	–
Stainless-steel fasteners	14%	72%	323	5
Total	19.75% (mean)	53.8% (mean)	3932	55

[a] (Import tons of alleged dumping country(ies)/Total extra EU imports in tons of product).
[b] The number of firms refers to the actual numbers used in the regression analysis. The initial number of firms retrieved was larger, but we lost firms due to missing observations on some of the variables needed in the analysis.
[c] Price undertakings. For the trade-weighted duty in this case, we used 4% for all firms.

Our empirical specification can be written as follows:

$$\Delta Y_{it} = \alpha_i + \mu_1 \Delta X_{it} + \mu_2[\Delta X_{it} \times AD] + \mu_3[\Delta X_{it} \times GDP_{kt}] + \beta_1 AD + \beta_2 GDP_{kt} + \varepsilon_{it}, \tag{7}$$

where μ_1 is the markup before protection, while μ_2 is the *change* in the markup during AD protection, which is our main interest. The total markup during protection is equal to $\mu_1 + \mu_2$. The change in the markup ratio due to business cycle fluctuations is captured by μ_3; α_i is a firm level fixed effect to capture firm heterogeneity; β_1 and β_2 measure the direct impact of the control variables AD protection and GDP growth in country k, and ε_{it} is a white noise error term that is added. In the remainder of our analysis, it has to be kept in mind that while our empirical specification allows us to estimate markups $(P/c)it = \mu_{it}$, it does not allow us to separate the price effect from the cost effect on the markup.

2.2. Data

The accounts data we use come from a commercial database sold under the name of AMADEUS that runs from 1992 to 2000. To identify the EU firms in the import competing industry, we collected information published in the Official EU Journal about the cases. In 1996, 26 new AD Investigations[13] were initiated, representing 12 different products or product groups. A product is very narrowly defined at the eight-digit Combined Nomenclature (CN) product classification used by the European Union. A

[13] A case here refers to a separate investigation per dumping country involved.

unique feature of our approach lies exactly in 'matching' these eight-digit products mentioned in the AD case, with the EU firms producing these products.[14]

In Table 1, we list the 10 product groups for which we could retrieve all the variables from the unconsolidated company accounts, required for our analysis. In column 2 of Table 1, we list the trade-weighted duty in the 'Protection' cases that was obtained by multiplying the duty per country by the import shares of the individual dumping countries in the EU. In six cases, the outcome was protection in the form of an AD duty. In one other protection case, 'Farmed Atlantic salmon', the outcome was 'mixed' with some importers subject to a duty and others subject to a price undertaking. Three cases out of the 10 were terminated without protection.

In the last columns of Table 1, we report the total number of EU firms we used for our analysis and the number of EU firms involved in the initiation of the filing of the complaint to the EU. For clarification, we point out that when the EU Commission decides to impose a duty, it applies to all EU-member states and can be compared to a 'common tariff' protecting the EU import competing industry as a whole and not just the EU firms that initiated the AD complaint.

3. Results

3.1. Pooled cases

3.1.1. Basic specification

In the first row of Table 2, we show the results of estimating Eq. (7) for the 'Protection' cases, where we pool all cases together. The results are obtained using a fixed effects specification that controls for unobserved firm heterogeneity.[15] The average markup, μ_1, before the filing decision in the 'Protection' cases lies around 16%, while the increase in markup after protection, μ_2, is 8% and significant at the 1% level.[16] We compare this increase in markups with two counterfactual samples. First, a number of AD complaints resulted in terminations. In row 2 of Table 2, we show the results for the pooled 'Termination' cases.[17] The average markup before filing lies around 26%, but we fail to find an increase in the markups after 1997.

As a second counterfactual, we randomly sample a control group of EU firms constraining the sampling to 10 sectors, different from the ones already in our data. In order to have a sufficient number of observations in each product group, we sampled sectors at the four-digit NACE level. In view of the high correlation between AD filings

[14] For full details on the data collection see Konings and Vandenbussche (2004).

[15] A Hausman test favored fixed effects over random effects, but results did not differ much. An F test for fixed effects was significant in all our specifications. OLS and robust regression yielded the same qualitative results, but will not be reported for brevity.

[16] The estimates for μ_3 were all statistically insignificant for the pooled cases and will not be reported for brevity.

[17] The large number of observations for the Termination cases is due to the more broadly defined products while in the Protection cases products were more specific.

Table 2
Estimation results (fixed effects)

$$\Delta Y_{it} = \alpha_i + \mu_1 \Delta X_{it} + \mu_2[\Delta X_{it} \times \mathrm{AD}] + \mu_3[\Delta X_{it} \times \mathrm{GDP}_{kt}] + \beta_1 \mathrm{AD} + \beta_2 \mathrm{GDP}_{kt} + \varepsilon_{it}$$

	P/c before AD protection (2)		Change in *P/c* during AD protection (3)		Number of observations (4)	R^2 (5)
	μ_1	S.E.	μ_2	S.E.		
Protection cases pooled	1.163***	(0.010)	0.079***	(0.013)	8708	0.83
Termination cases pooled	1.257***	(0.012)	0.011	(0.015)	7214	0.80
Counterfactual pooled	1.213***	(0.007)	− 0.006	(0.010)	15,591	0.85
Protection cases correcting for selection bias	1.178***	(0.030)	0.087**	(0.039)	8708	–
(Inverse mills ratio)	0.002	(0.003)				
Protection cases separately						
Cotton fabrics	1.27***	(0.033)	0.065*	(0.044)	777	0.86
Leather handbags	1.237***	(0.011)	0.055***	(0.014)	5045	0.89
Farmed Atlantic salmon	1.039	(0.030)	0.129***	(0.038)	1710	0.75
Seamless steel pipes and tubes	0.993	(0.033)	− 0.049	(0.047)	695	0.75
Polyester fibre yarns	1.088*	(0.050)	0.191***	(0.075)	528	0.72
Bed linen	1.61***	(0.064)	0.36***	(0.101)	151	0.90
Stainless-steel fasteners	1.17***	(0.019)	0.091***	(0.023)	1610	0.91
Termination cases separately						
Synthetic fibre ropes	1.19***	(0.026)	0.033	(0.034)	883	0.88
Luggage and travel goods	1.27***	(0.013)	0.007	(0.017)	6262	0.83
Video tapes	1.38***	(0.175)	0.080	(0.24)	101	0.63
Counterfactual cases separately						
Plastics in primary form	1.203***	(0.022)	0.033	(0.031)	218	0.86
Metal structures	1.224***	(0.023)	− 0.031	(0.030)	2228	0.80
Copper	1.217***	(0.039)	− 0.105**	(0.050)	637	0.83
Processing of meat	1.131***	(0.008)	0.000	(0.012)	2251	0.95
Processing of fruit and vegetables	1.175***	(0.026)	− 0.005	(0.035)	1459	0.80
Grain mill products	1.132***	(0.020)	− 0.016	(0.028)	1033	0.89
Wine	1.153***	(0.026)	− 0.021	(0.032)	1470	0.84
Outwear	1.338***	(0.019)	0.020	(0.026)	3481	0.79
Inorganic basic chemicals	1.223***	(0.042)	0.067	(0.056)	659	0.77
Cement	1.344***	(0.044)	− 0.074	(0.053)	978	0.80

(i) The coefficient μ_1 is the markup in the absence of protection, and we test whether it is statistically different from 1. The parameter that captures the *change* in market power from 1997 onwards is given by μ_2. We test for it to be statistically different from zero. Standard errors in brackets; ***/**denotes statistically significant at the 1%/5% critical level or lower. (ii) The mean growth of GDP over the period was 2%. (iii) The estimated coefficients of the interaction term with GDP growth (μ_3) and of the extra control variables (β_1, β_2) were not statistically significant and we therefore did not report them for brevity. (iv) An F test indicated that the fixed effects turned out to be statistically significant at the 5% critical level or lower.

and ‘openness’ of a sector, we constrained the random sampling of our control group in the 25% most open sectors, clustered around 10 different product groups, but excluding those sectors that had been subject to AD filings in the past. The results of our basic specification in Eq. (7) on the pooled sample of these randomly selected product groups are shown in the third row of Table 2. The μ_1 coefficient, giving the level of the markup before 1996 for the total control group of firms, is 21%. But the μ_2 coefficient capturing the change in markup after 1997 is negative and insignificant. This suggests that for the firms in our control group we do not find an increase in markups after 1997.

Two additional results are worth pointing out here (for a more detailed discussion see Konings and Vandenbusssche, 2004). First, when we run our basic Eq. (7) separately for single product firms and for multi-product firms, we find that the markups of single product firms protected by AD duties increased by 14 percentage points, while in the case of multi-product firms the change was only 7 percentage points. Since most of the firms in our sample are multi-product firms, our results reported earlier are likely to be a lower bound of the true rise in markups at the product line. Second, when running Eq. (7) with an additional interaction dummy for EU firms involved in the filing of the dumping complaint, we find on average an additional 19 percentage points bonus on markups. However, this result should be treated with the necessary caution in view of the small number of observations on initiating firms as indicated in Table 1.

3.1.2. Using AD duty levels

In Table 2, we used an AD dummy, without taking into account the magnitude of the duty. In the first column of Table 3, we replace the AD dummy with the trade-weighted AD duties in the Protection cases.[18] Averaging the trade-weighted duties over all cases gave us a duty level for the 96 Protection cases of about 20% with a standard deviation of 12%. This may seem low compared to estimates for the US, where the average dumping margin is around 65% (Blonigen, 2003). One of the reasons is the ‘lesser-duty-rule’ that prevails in the EU but not in the US, which implies that the AD duty can be smaller than the dumping margin.

The results in the first column of Table 3 show that the average markup across products is about 16%, comparable to what we had in Table 2 using the AD dummy. The change in the markup coefficient is equal to 0.302 and is statistically significant. The effect of the duty evaluated at the average duty level is $0.302 \times 0.20 = 0.061$ or about 6 percentage points, which compares quite well with the 8 percentage point increase obtained with the AD dummy. Note also that in Table 3, we now have a negative and significant effect of GDP growth on markups. The effect on markups evaluated at the mean GDP growth of 2% in our sample amounts to a reduction in markups of less than 1 percentage points.

3.1.3. Import diversion effects

In order to check to what extent markups vary with imports of dumping countries, we include in our basic specification the ‘log of imports in tons of the dumping countries’ instead of the AD variable.[19] We expect a negative sign for this interaction term in the sense that when imports of the dumping countries fall after protection (Prusa, 1997), we expect to see domestic markups go up. To capture the idea of trade diversion, we also include in our basic specification an additional interaction term with the ‘log of imports of the non-dumping countries’.[20] Here, we also expect a negative sign of the coefficient of

[18] In the ‘mixed’ case we assumed the price undertaking to be equal to the country wide duty level.

[19] We do not have information on the imports of individual dumping firms, therefore, we take as a proxy the product level imports from dumping and non-dumping countries similar to what Prusa (1997) used.

[20] Data were obtained from EUROSTAT, annual intra- and extra-EU imports (2001).

Table 3
Estimation results for pooled cases (fixed effects), using AD duty levels

	Protection cases (1)	Protection cases (2)	Protection cases (using lagged imports $t-2$) (3)
	$\mu_1 = P/c$	μ_1	μ_1
ΔX (= composite explanatory variable of nominal inputs weighted by factor shares)	1.163*** (0.008)	1.59*** (0.066)	1.74*** (0.09)
	$\mu_2 = \Delta P/c$	μ_2	μ_2
ΔX interacted with AD trade-weighted duty level	0.302*** (0.037)	–	–
ΔX interacted with log of dumping imports	–	− 0.028*** (0.004)	− 0.036*** (0.006)
ΔX interacted with log of non-dumping imports	–	− 0.013 (0.010)	− 0.018 (0.013)
ΔX interacted with annual country level GDP growth	− 0.101* (0.058)	− 0.198*** (0.058)	− 0.30*** (0.077)
R^2	0.83	0.83	0.81
Number of observations	8708	8708	5934

Similar to Table 2. The AD duty, GDP growth and the log of imports in tons from dumping countries and from non-dumping countries have also been included in the regressions separately as extra controls in addition to their interactions but results are not reported for brevity.

the interaction term. If trade diversion takes place, the non-dumping countries start exporting more to the EU market, putting downward pressure on the domestic markups.

The results are reported in column (2) of Table 3. The 'log of imports of dumping countries' has a negative and significant effect on markups, while the 'log of imports of the non-dumping countries' has a negative and insignificant effect on markups, suggesting that trade diversion is far from complete in the AD cases we consider. However, the inclusion of imports may result in an endogeneity bias, therefore, in the final column of Table 3, we also run a specification where we lag imports of the dumping and of the non-dumping countries by 2 years. The results do not change. The dominant effect on markups is the imports of the dumping countries.

3.2. Individual cases

We now turn to the 10 different products in which the AD investigations took place. The markups and the change in markups are presented in Table 2. In all but one of the affirmative AD cases, we find the change in the average European markup as a result of AD protection to be positive and significant, with markup increases ranging between 5.5 percentage points in the case of 'Leather handbags' and 36 percentage points in the case of 'Bed linen'. The exception to the general increase in markups due to AD protection is 'Seamless steel pipes and tubes' where we do not find any effect of AD protection on the average markup, despite the relatively high trade-weighted duty level of 27.9% applying in this case. The explanation seems to lie partly in the substantial trade diversion that followed the protection decision, as illustrated in Fig. 1. While dumped imports fell after

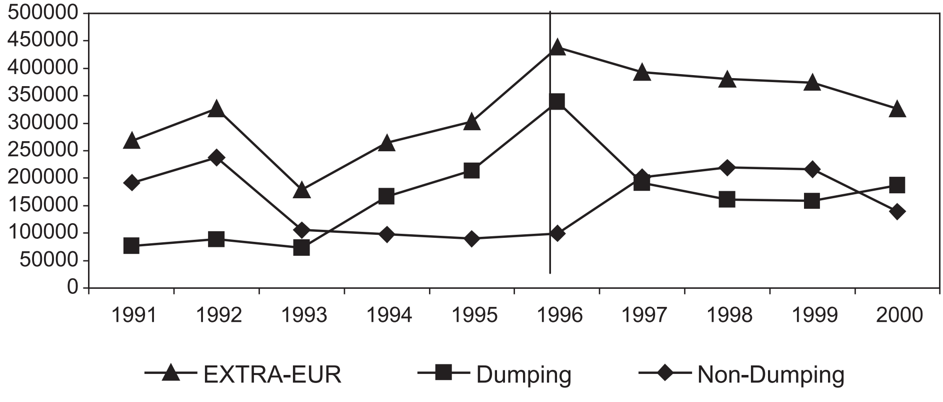

Fig. 1. Evolution of imports in metric tons of ‘Seamless steel tubes case’.

1996, there was a simultaneous increase in imports from the non-dumping countries, leaving the total extra-EU imports of ‘Seamless steel pipes and tubes’ quite stable after protection. This substantial trade diversion is likely to have prevented the domestic EU market for ‘Seamless steel pipes and tubes’ from raising their markups.

Quite a different pattern emerges in the ‘Leather handbags’ case as illustrated in Fig. 2. While imports of the dumping countries fell after 1996, the imports of the non-dumping countries remained at the same level as before the protection. There was clearly far less trade diversion going on, if any. As a result, less total imports entered the EU market after 1996, presumably allowing domestic EU markups to rise by 5.5 percentage points over and above the 23% markup that already prevailed in the market before protection.

For the three Termination cases, ‘Synthetic fibre ropes’, ‘Luggage and travel goods’ and ‘Video tapes’, we fail to find a significant increase in markups after 1996. This seems to suggest that changes in markups are driven by the outcome of a case and not so much by the AD filing decision, as already suggested by the pooled results.

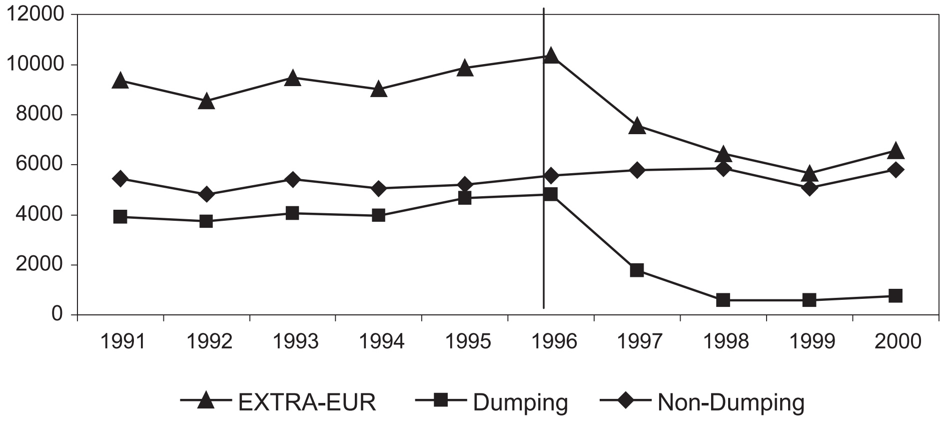

Fig. 2. Evolution of imports in metric tons of ‘Leather handbags’.

Finally, we turn to a control group of 10 different product groups in the EU, listed in the bottom half of Table 2. Markups before 1996 in these sectors range from 13% to 34%. However, none of the product groups in our EU control group experienced a significant increase in markups after 1996.

4. Robustness checks

4.1. Controlling for selection bias

One important concern we need to address is that the filing of an AD case may depend on certain firm and industry characteristics.[21] The pattern of markups before protection speaks to this issue (Fig. 3). For ‘Protection cases’, the yearly evolution of markups before protection is quite erratic up to 1996. After protection sets in, there seems to be an upward trend in markups. This suggests that it is unlikely that the average increase in markups that we find in the ‘Protection cases’ is due to a self-selection of firms with rising/declining markups receiving AD protection. The evolution of markups in the ‘Terminations’ is much more stable over time but without a clear upward trend afterwards. Finally, for the firms in the ‘Control group’, a pattern emerges that is more similar to the ‘Termination cases’ than to the ‘Protection cases’.

We test more formally for the presence of selection bias by using a two-step Heckman procedure (Greene, 2000, p. 930). We used all AD filings aggregated up to the four-digit sector level between 1995 and 2000 to estimate the probability of AD filing as a function of import penetration, industry employment, sales growth and previous AD filings, as in Blonigen and Park (2001).[22] From this equation, we computed the inverse Mills ratio and included it in Eq. (7). Given that the computed Mills ratio resulted from a separate estimation, we adjusted the standard errors in the main analysis by bootstrapping it a 1000 times (Hill et al., in press). The results are reported in row (4) of Table 2. We note that for the pooled sample of firms in the Protection cases, the price–cost markup is now estimated at about 18% and its increase after AD protection is estimated at almost 9 percentage points, which is very similar to the results under the fixed effects specification in row (1). Furthermore, the inverse Mills ratio is not statistically significant different from zero, which suggests that our basic specification is not subject to selection bias.

4.2. Endogeneity

A second concern is that firms filing for protection may expect their prices to go up in the future which can affect input demands. This can be seen from Eq. (6) where output prices occur on the left-hand side and inputs on the right-hand side of the equation. This

[21] For example, Staiger and Wolak (1994), Knetter and Prusa (2000) and Blonigen and Park (2001).

[22] The results of this selection equation are omitted for brevity, but can be found in Konings and Vandenbussche (2004).

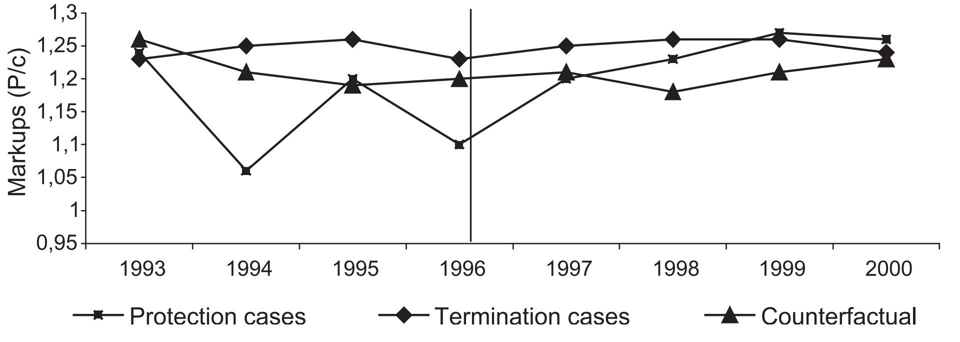

Fig. 3. Evolution of markups for the pooled 1996 antidumping cases.

may result in endogeneity of ΔX in Eq. (7). Similarly, endogeneity may come from measurement error in the input variables.

To account for these two sources of endogeneity, we estimate Eq. (7) with instrumental variables using the general methods of moments estimator (GMM) proposed by Arellano and Bond (1991) (AB). This implies that we use as instruments all lagged values of ΔX starting from $t-2$ and before and estimate Eq. (7) in first differences to control for unobserved fixed effects. Table 4 shows the results. While the point estimates are quite different compared to those reported in Table 2,[23] we continue to find a significant increase in markups in the 'Protection' cases, while no statistically significant increase in the 'Termination' and 'Counterfactual' cases was found. We note that the Sargan test confirms the instrument validity in all cases and that the second order serial correlation test (SOC) does not reject the model.

Another concern is that the methodology we used so far did not control for any dynamics in the markups. If AD protection is induced by low markups, and if markups are mean-reverting, the observed post-duty increases in markups might have little to do with the effects of duties on market power.

To explore this possibility, we turn to an alternative approach to measuring market power. As discussed by Tybout (2003), a common approach is to use the *observed firm level price–cost margin* (PCM), defined as sales net of expenditures on labor and materials over sales ($\mathrm{PCM}_{it}=(P_{it}Q_{it}-P_{Mit}M_{it}-P_{Nit}N_{it})/P_{it}Q_{it}$). We follow the literature and specify the following regression equation:

$$\mathrm{PCM}_{it}=\gamma_i+\gamma_1\mathrm{PCM}_{it-1}+\gamma_2(K_{it}/P_{it}Q_{it})+\gamma_3\mathrm{AD}+\gamma_4\mathrm{GDP}_{kt}+\varphi_{it}, \tag{8}$$

where γ_i is an unobserved firm level fixed effect and φ_{it} is a white noise error term. The lagged dependent variable is included to control for the possibility that price–cost margins are mean-reverting. As additional controls we include the capital–sales ratio, GDP growth in country k at time t, year dummies and case dummies. We estimate Eq. (8) in first differences using GMM and instrument the lagged dependent variable PCM_{it-1} and the

[23] This may be explained by the choice of instruments, i.e., lagged values of ΔX.

Table 4
Estimation results using IV Arellano–Bond GMM estimator

$$\Delta Y_{it} = \alpha_i + \mu_1 \Delta X_{it} + \mu_2 [\Delta X_{it} \times \mathrm{AD}] + \mu_3 [\Delta X_{it} \times \mathrm{GDP}_{kt}] + \beta_1 \mathrm{AD} + \beta_2 \mathrm{GDP}_{kt} + \varepsilon_{it}$$

	Protection cases (1)	Termination cases (2)	Counterfactual (3)
	$\mu_1 = P/c$	$\mu_1 = P/c$	$\mu_1 = P/c$
ΔX (=composite explanatory variable of nominal inputs weighted by factor shares)	1.05*** (0.11)	1.06*** (0.16)	1.213*** (0.18)
	$\mu_2 = \Delta P/c$	$\mu_2 = \Delta P/c$	$\mu_2 = \Delta P/c$
ΔX interacted with AD dummy (1 from 1997 onwards)	0.21** (0.13)	0.17 (0.20)	− 0.33 (0.26)
	μ_3	μ_3	μ_3
ΔX interacted with annual country level GDP growth	1.12 (0.7)	− 1.03 (0.91)	− 2.8** (1.60)
Sargan test of instrument validity (p value)[a]	0.16	0.40	0.19
Test of second-order serial correlation[b]	1.35	1.65	1.25
Year dummies	Yes	Yes	yes
Number of observations	4468	3368	9645

***/* Denotes statistically different from zero at the 1%/5% critical level. Instruments include all available moment restrictions of ΔX starting at $t-2$ and before.

[a] Which asymptotically follows a χ^2 distribution.

[b] Which asymptotically follows a $N(0,1)$ distribution.

capital intensity variable with their lagged values dated $t-2$ and before, as they are not correlated with the first differenced error term.

The results are shown in Table 5. The point estimates suggest that the firm level PCM is on average 4 percentage points higher after protection, while we find no significant increase in the firm level PCM in the 'Termination' cases or the 'Counterfactual'. Given that the sample average PCM in the Protection cases is 27% before protection, an increase of 4 percentage points in the PCM after protection is equivalent to an increase of 7.5

Table 5
PCM method: first differences GMM estimates

	Protection cases	Termination cases	Counterfactual
PCM_{t-1}	0.233*** (0.03)	0.215*** (0.04)	0.242*** (0.039)
Capital intensity	− 0.04*** (0.01)	0.022 (0.026)	− 0.035** (0.016)
AD protection	0.04** (0.02)	0.001 (0.016)	− 0.006 (0.016)
GDP growth	0.13*** (0.03)	0.098*** (0.03)	0.068*** (0.015)
Year dummies	Yes	Yes	Yes
Case dummies	Yes	Yes	Yes
Constant	− 0.02*** (0.007)	− 0.012*** (0.008)	− 0.02*** (0.004)
Sargan test	0.51	0.60	0.30
Test second-order serial correlation	0.32	0.031	1.9
Number of observations	4468	3368	9645

***/** Denotes statistically significant different from zero at the 1%/5% level. Instruments include the moment restrictions on PCM and capital intensity at $t-2$ and before.

percentage points in the markup (P/c). This is very similar to our basic estimates reported in Table 2.

Thus, irrespective of the method used, we find evidence of a positive effect on firm markups after AD protection.

5. Conclusion

In this paper, we report evidence that markups of domestic firms increase significantly during AD protection against dumped imports. In view of the firm level data we use, our results can be considered as a lower bound of the true rise in markups at the product line. AD filings without ensuing protection did not result in increased markups, suggesting that the 'protection' decision rather than the 'filing' decision is required for rising markups. Import diversion has a negative effect on markups and is more prevalent in some industries than in others.

It is not clear a priori whether the same results would hold for the US. The absence of the 'lesser-duty' rule and the presence of an 'Administrative Review process' in the US (Blonigen and Haynes, 2002) result in higher duties than in the EU. However, the substantial trade diversion reported for the US by Prusa (1994), could discipline markups more than in the EU, where trade diversion seems less strong (Konings et al., 1999). Hence, the effect of AD duties on markups for the US is difficult to predict, and remains an interesting topic for future research.

Acknowledgements

We thank A. Bernard, B. Blonigen, C. Brown, E. Edmonds, D. Irwin, J. Harrigan, K. Krishna, N. Pavcnik, T. Prusa, D. Richardson, W. Roeger, J. Van Biesebroeck, F. Verboven and F. Warzynxki for comments and discussions. We also thank two anonymous referees for numerous suggestions and the co-editor Jim Tybout for additional guidance.

References

Arellano, M., Bond, S., 1991. Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations. Review of Economic Studies 58 (2), 277–297.

Belderbos, R., Vandenbussche, H., Veugelers, R., 2004. Antidumping jumping FDI in the EU. European Economic Review 48 (2), 429–452.

Blonigen, B., 2003. Evolving discretionary practices of U.S. antidumping activity. NBER Working Paper 9625, 36.

Blonigen, B., Haynes, S., 2002. Antidumping investigations and the pass-through of AD duties and exchange rates. American Economic Review 92 (4), 1044–1061.

Blonigen, B., Park, J.-H., 2001. Dynamic pricing in the presence of antidumping-policy: theory and evidence. NBER Working Paper 8477, 49.

Blonigen, B., Prusa, T., 2003. Antidumping. In: Choi, K., Harrigan, J. (Eds.), Handbook of International Trade. Basil Blackell, Oxford, UK and Cambridge, MA, pp. 456–476.

Fischer, R., 1992. Endogenous probability of protection and firm behavior. Journal of International Economics 32, 149–163.

Greene, W., 2000. Econometric Analysis, 4th ed. Prentice Hall International Edition, New Jersey, US.

Hall, R.E., 1988. The relation between price and marginal cost in the US industry. Journal of Political Economy 96, 921–947.

Hall, R.E., Jorgenson, D.W., 1967. Tax policy and investment behavior. American Economic Review 57, 391–414.

Harrison, A.E., 1994. Productivity, imperfect competition and trade reform. Journal of International Economics 36, 53–73.

Helpman, E., Krugman, P., 1989. Trade Policy and Market Structure. MIT Press, Cambridge.

Hill, R.C., Adkins, L.C., Bender, K.A., 2003. Test statistics and critical values in selectivity models. In: Carter Hill, R., Fomby, T.B. (Eds.), Advances in econometrics: maximum likelihood estimation of misspecified models—twenty years later 17. Elsevier Science. forthcoming.

Hsieh, C.T., 2002. What explains the industrial revolution in East-Asia? Evidence from factor markets. American Economic Review 92 (3), 502–526.

Knetter, M., Prusa, T., 2000. Macroeconomic factors and antidumping filings: evidence from four countries. NBER Working Paper 8010, 42.

Konings, J., Vandenbusssche, H., 2004. Antidumping Protection and Markups of Domestic Firms. LICOS discussion paper 141, Catholic University of Leuven http://www.econ.kuleuven.ac.be/licos/DP/DP2004/DP141.pdf), p. 42.

Konings, J., Vandenbussche, H., Springael, L., 1999. Import diversion as a result of European antidumping policy. NBER Working Paper 7340, 39.

Krishna, P., Mitra, D., 1998. Trade liberalization, market discipline and productivity growth: new evidence from India. Journal of Development Economics 56 (2), 447–462.

Lenway, S., Rehbein, K., Starks, L., 1990. The impact of protection on firm wealth: the experience of the steel industry. Southern Economic Journal 56, 1079–1093.

Levinsohn, J., 1993. Testing the imports-as-market-discipline hypothesis. Journal of International Economics 35, 1–22.

Olley, G., Pakes, A., 1996. The dynamics of productivity in the telecommunications equipment industry. Econometrica 64 (6), 1263–1297.

Pauwels, W., Vandenbussche, H., Weverbergh, M., 2001. Strategic behavior under European antidumping policy. International Journal of the Economics of Business 8, 79–103.

Pavcnik, N., 2002. Trade liberalization, exit and productivity improvements: evidence from Chilean plants. Review of Economic Studies 69, 245–276.

Prusa, T., 1994. Pricing behavior in the presence of AD law. Journal of Economic Integration 9 (2), 260–289.

Prusa, T., 1997. The trade effects of U.S. antidumping actions. In: Feenstra, R. (Ed.), The Effects of US Trade Protection and Promotion Policies—NBER Volume. The University of Chicago Press, pp. 191–214.

Reitzes, J., 1993. Antidumping policy. International Economic Review 34 (4), 475–763.

Roeger, W., 1995. Can imperfect competition explain the difference between primal and dual productivity measures? Estimates from US manufacturing. Journal of Political Economy 103, 316–330.

Rotemberg, J., Woodford, M., 1991. Mark-ups and the business cycle. In: Blanchard, O.J., Fischer, S. (Eds.), NBER Macroeconomics Annual 1991. MIT-Press, Cambridge, MA, p. 306.

Staiger, R., Wolak, F., 1994. Measuring industry specific protection: antidumping in the U.S. Brookings Papers on Economic Activity. Microeconomics 1, 51–118.

Tybout, J., 2003. Plant- and firm-level evidence on "New" trade theories. In: Choi, K., Harrigan, J. (Eds.), Handbook of International Trade. Basil Blackwell, UK, Oxford and US, MA.

Veugelers, R., Vandenbussche, H., 1999. European antidumping policy and the profitability of national versus international collusion. European Economic Review 47 (1), 1–25.

Zanardi, M., 2004. Antidumping law as a collusive device. Canadian Journal of Economics 37 (1), 95–122 (February).