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Macroeconomic Determinants of Contingent Protection:

The Case of the European Union

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Abstract:

Contingent Protection has grown to become an important trade restricting device. In the European Union, protection instruments like antidumping are used extensively. This paper analyses whether macroeconomic pressures may contribute to explain the variations in the intensity of antidumping protectionism in the EU. The empirical analysis uses count data models, applying various specification tests to derive the most appropriate specification. Our results suggest that the filing activity is inversely related to the macroeconomic conditions. Moreover, they confirm existing evidence for the US suggesting that domestic macroeconomic pressures are a more important determinant of contingent protection policy than external pressures.

JEL Classification: F13, F4

Keywords: Antidumping; Protection; Macroeconomic Conditions; Estimation of Count Data Models

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I. Introduction

Contingent protection (CP) measures are GATT legal tools of protection. The most important instruments of contingent protection are safeguard measures (article XIX GATT-1994) as well as anti-dumping (AD) measures and countervailing (CV) measures, both based on article VI GATT-1994. According to *Finger (1993)*, the GATT recorded more than 2000 CP cases during the 1980s, mostly initiated by the United States, the EU[1], Canada and Australia. More recently, contingent protection has evolved into a global phenomenon as more and more transition and developing countries established CP-laws and started to make use of these (*Miranda and Torres, 1997*).

The growing importance of contingent protection raises the question of its determinants. According to the GATT rules, actions are contingent on industry-specific circumstances: safeguard actions can only be adopted if an increase in imports has caused (or threatens to cause) serious injury to the domestic industry. In case of AD or CV actions, the importing country must demonstrate that imports are dumped and consequently materially injure the domestic industry. Therefore, most studies have adopted an industry-specific perspective to explain the use of contingent protection, while only few studies have investigated whether contingent protection actions may be related to macroeconomic conditions.[2]

This paper is the first to analyse the macroeconomic determinants of contingent protection policy for the European Union[3]. Unlike previous studies for the US, who have mostly relied on OLS estimation techniques, count data models are employed in order to account for the discrete and non-negative nature of the data generating process. Various specification tests are conducted to derive the most appropriate specification. The plan of the paper is as follows: *Section II* briefly describes the institutional set up and selected stylised facts of contingent protection in the EU. *Section III* reviews the relevant literature. In *Section IV*, we present and discuss the model and our empirical findings. *Section V* concludes.

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1 Although the three European Communities (EC, ECSC and EAEC) technically still exist, and the European Union as such does not have a legal personality, throughout this paper, the term European Union or EU shall be used to denote the Communities.

2 For a discussion of the industry-specific studies, see *Tharakan (1995)*.

3 The analysis is done for the period 1980-98. The year 1980 is the first year covered by an annual report of the Commission of the European Communities on the Community’s antidumping and antisubsidy activities.

II. Contingent Protection Policy in the European Communities: Institutional Set Up and Stylised Facts

In the EU, the bulk of contingent protection policy falls on the instrument of antidumping. Between 1980 and 1998, antidumping cases accounted for almost 95 percent of all European contingent protection cases. 669 antidumping cases were launched as opposed to only 19 (10) countervailing (escape clause) investigations[4]. Hence, the contingent protection policy in the EU is predominantly an antidumping policy.

AD-measures serve to protect domestic firms from "unfair" foreign import competition. According to European AD-legislation[5], which is in accordance with the WTO AD-agreement, dumping is usually defined as international price discrimination, i.e. dumping is given, if the import price is below the domestic market price of a certain "like" product. European trade legislation allows to impose AD duties on imports if the European AD authority, the European Commission, proves that dumping has occurred and has caused injury to the domestic industry. Alternatively, in case of an affirmative dumping and injury finding, the Commission also has the right to terminate proceedings by undertakings: in this case, the exporters "voluntarily" increase their prices to an extent which removes the injurious effects of dumping.

AD investigations are requested by the domestic industry. The Commission decides whether a formal complaint contains sufficient prima facie evidence justifying the initiation of an investigation. During 1980-1998, this has been the case in – on average - 36 cases per year, which reflects presumably both: firms that are active in filing complaints as well as an AD authority who accepts the complaints leniently.[6] Approximately two thirds of these cases end with affirmative dumping and injury findings (*Vandenbussche, Konings and Springael, 1999*). Hence, once a case has been opened, domestic firms face a high probability of obtaining protection via AD duties or undertakings. Moreover, theoretical (*Prusa, 1992; Panagariya and Gupta, 1998*) and empirical work (*Messerlin, 1989; Staiger and Wolak, 1994*) has found that imports may fall even if no dumping and injury is found, since cases are frequently

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4 See Annual Reports of the European Commission to the European Parliament, various issues.

5 Since 1979, there have been five different antidumping regulations (Regulation No. 1681/79, 2176/84; 2423/88, 3283/94; 394/96. The following description holds for all of these regulations.

6 The Commission withholds information on this pre-investigation stage. The number of complaints rejected by the Commission on grounds of missing evidence for dumping is thus not available.

withdrawn by the complainants as a reaction to an out-of-court settlement between the domestic industry and the accused exporters. In this settlement the foreign competitors commit themselves to increase prices and reduce exports. Because of such investigation or withdrawal effects, the number of investigations initiated each year is usually regarded to be a better proxy for the protective effect than the number of cases that end with affirmative decisions on injurious dumping (see e.g. *Leidy, 1997*). In the EU, the case activity per year is subject to considerable variation: The number of newly initiated cases frequently jumps after an external shock before protectionist pressure gradually declines in the subsequent periods. In some years, the case activity was relatively low, i.e. below 25 cases per year, while in others, it increased to more than 50 or - as most recently (1999) - to 86 cases.

III. Macroeconomic Determinants of Contingent Protection

The observed variations in the case activity of EU antidumping policy may be related to macroeconomic determinants. Hereby, two different channels can be distinguished: First, the balance of payment situation may have an impact on the willingness to accept a complaint if the national policy makers pressure the Commission to use trade protection as a tool of expenditure switching. According to this *external pressure hypothesis,* the number of antidumping cases per year is therefore positively related to a widening in the trade balance deficit or to a real appreciation of the domestic currency.[7]

Second, the domestic macroeconomic situation may influence the filing activity of domestic firms: if the domestic macroeconomic activity is sluggish, and unemployment relatively high, any further increase in import competition puts downward pressure on each worker’s wage in case he is dismissed. This tends to increase lobbying efforts by unions. Additionally, rent seekers may anticipate that the governments are sensitive to any further increase in imports which threaten to cause layoffs. According to this *domestic pressure hypothesis*, rent seeking pressures increase in recessions and vent by dumping complaints.

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[7] Exchange rate swings may also matter in antidumping cases by inducing pricing-to-market behaviour (Feinberg, 1989). Pricing to market occurs when firms do not pass through nominal exchange rate swings into their export prices. When the exporting country’s currency is appreciating, WTO rules induce the AD-authority to interpret pricing-to-market by foreign firms as dumping, since foreign export prices expressed in foreign currency are then lower than in their domestic markets.

Table 1: Compilation of Previous Work on the Macroeconomic Determinants of Contingent Protection

Author	*Case*	*Explanatory Variable*[a]		*Significant Determinants of Contingent Protection*
		Domestic Pressure	*External Pressure*	
Takacs (1981)	Escape clause	Unemployment rate, capacity utilisation rate, level of nominal GNP	Trade balance	Domestic and external pressure
Feigenbaum and Willett (1985)	Escape clause	Capacity utilisation rate	Trade balance, import penetration, balance of goods and services, consumer price index, real exchange rate	Domestic pressure only
Salvatore (1987)	Escape clause	Level of real GNP, unemployment rate, capacity utilisation rate	Trade balance	Domestic pressure only
Coughlin, Terza and Kahalifah (1989)	Escape clause	Capacity utilisation rate	Trade balance	Domestic and external pressure
Leidy (1997)	Antidumping and countervailing Cases	Unemployment rate, capacity utilisation rate	Trade balance, import penetration, real effective exchange rate	Domestic pressure only

[a] Concentrating on the proxies for domestic and external pressure.

Existing empirical evidence, focussing on the US experience, has confirmed that macroeconomic pressure has an influence on the course of contingent protection policy over time (see table 1). Hereby, all studies find that the case activity is related to internal pressure variables approximated by changes in the rate of capacity utilisation and/or the unemployment rate as well as in the level of GNP[8]. The studies however differ in their evaluation of whether external pressure matters or not: evidence in favour of the external pressure hypothesis was provided by *Takacs (1981)* and *Coughlin, Terza and Khalifah (1987)*, while evidence indicating their unimportance follows from the work of *Feigenbaum and Willet (1985), Salvatore (1987)* and *Leidy (1997)*. Regarding the estimation tools employed in the different studies, they in most part relied on conventional OLS regression techniques. However, as mentioned in the introduction and further explained below, while they might provide a reasonable approximation for large counts, they cannot capture the discrete and non-negative nature of count data. *Coughlin, Terza and Khalifah (1987)* ran their regressions using a Poisson, i.e. a count model, a Box Cox and an OLS specification, showing that the Poisson specification dominates the other two. For these reasons, we employed various count data specifications for our analysis.

IV. Empirical Results

1. Model Specification

In order to analyze macroeconomic influences on the pressure for antidumping protection, we use the following econometric model:

$$NUMBERNEW(t) = f\left[IM(t-1), EM(t-1), IPJAPGR(t-1), NUMBERNEW(t-1)\right] \qquad (1),$$

t=1,...,T, where NUMBERNEW(t) denotes the number of newly initiated AD and CV investigations per year t, IM(t-1) are internal macroeconomic pressures, EM(t-1) external macroeconomic pressures, and IPJAPGR(t-1) is the growth rate of total industrial production in Japan, all in year t-1.

In line with previous studies, our dependent variable is the number of investigations initiated per year (NUMBERNEW) rather than of those cases ending with affirmative decisions on dumping as the adequate measure for the intensity of contingent protectionist pressures. This is because of the investigation effect described in *section II*. Note once more that high values

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[8] The latter is used by *Takacs (1981)* and *Salvatore (1987)*. In a comment to Takacs' study, *Feigenbaum, Ortiz and Willet (1985)* rightly criticize using the level as a proxy for the cyclical condititon. Rather, the growth rate would serve as an appropriate proxy. Among others, we use the growth rate of real GDP in our study.

of the dependent variable may imply either a relatively lenient willingness to accept complaints by the Commission and/or a pronounced filing activity of the domestic industry in a given year. Regarding the independent variables, similar to studies for the US, factors exerting domestic or external macroeconomic pressure are distinguished between. *Domestic* macroeconomic pressure is approximated by the growth rate of real GDP (GDPGR(t-1)) and by that of the total industrial production (IPGR(t-1)) as well as the percentage change in the unemployment rate (UER(t-1)). We expect the coefficients of GDPGR(t-1) and of IPGR(t-1) to be negative and that of UER(t-1) to be positive. Indicators of external pressure are the real effective exchange rate (REER(t-1)), the trade balance (TB(t-1)), and the ratio of import penetration (IMPPEN(t-1), percentage change from previous period). Under the external pressure hypothesis, the coefficients of REER(t-1) and of TB(t-1) should be negative[9], while that of IMPPEN(t-1) should be positive. The variables and their description are summarized in table 2.

The growth rate of total industrial production in Japan (IPJAPGR(t-1)) was also included in the regressions as a proxy of the macroeconomic situation of the EU's trading partners. An economic downturn of a major trading partner may have an effect on the number of petitions filed either for reasons independent of the domestic macroeconomic conditions[10] or for those related to the balance of payments. In particular, in a recession, the exporter's domestic (here: the Japanese) markets absorb a considerably smaller share of supply, ceteris paribus increasing the export volume, i.e. the import competition faced by the trading partner (here: the EU) and hence, other things equal, increasing the trading partner's industries' demand for protection. Thus, theory suggests that IPJAPGR(t-1) should enter the regressions with a negative sign. The Japanese growth rate of total industrial production was chosen as a proxy for these outside influences as it was the most important target of European AD-policy during the investigation period, if measured by the trade volume affected. By the end of 1996, approximately one third of the total trade affected by AD-measures referred to imports from Japan.[11]

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9 The real effective exchange rate is defined so that an increase represents an improvement in the international competitive position.

10 For an elaboration of this idea see *Leidy (1997)*.

11 The number of cases initiated against Japanese firms during the period of investigation was 42.

Table 2: Dependent and Independent Variables used in the Regressions

Abbreviation	*Description*
<u>*Dependent variable*</u>	
NUMBERNEW(t)	Total number of newly initiated antidumping and antisubsidy cases per period
<u>*Independent variables*</u>	
Macroeconomic activity:	
GDPGR(t-1)	Growth rate of real gross domestic product per period
IPGR(t-1)	Growth rate of total industrial production per period
UER(t-1)	Unemployment rate per period (percentage change from previous period)
International trade position:	
REER(t-1)	Real effective exchange rate per period
TB(t-1)	Trade balance per period
IMPPEN(t-1)	Import penetration per period (= Imports/GDP, percentage change from previous period)
Further control variables:	
IPGRJAP(t-1)	Growth rate of total industrial production in Japan per period
NUMBERNEW(t-1)	Total number of newly initiated antidumping and antisubsidy cases per period (lagged dependent variable)

Note: The independent variables were lagged one period in order to account for the lagged effects of the variables on the economy and in order to avoid the possibility of reverse causation.

The lagged dependent variable NUMBERNEW(t-1) was included in order to model potential dependencies across time periods. A significantly negative sign would suggest a ‚depletion effect‘ à la *Leidy (1997)*: the higher the number of petitions is in a year t, the more the stock of potential petitions in the following year t+1 is depleted. According to Leidy, the depletion effect indicates the "safety value" nature of AD petitions. In case of macroeconomic downturns, protectionist pressure intensifies and is vented by AD petitions, which in turn, implies reduced demand for protection in the subsequent period. Hence, ceteris paribus, less petitions should be filed and also the acceptance rate of petitions in year t+1 may depend negatively on the number of cases opened in the preceding year.[12]

[12] *Leidy (1997)* suggests that this is due to the stock of petitioners being finite and the petitions remaining under consideration during the following year.

Finally, it was attempted to account for the introduction of the new antidumping regulation in the EU that came into force following the implementation of the Uruguay Round Agreements in 1995. In order to control for a possible “regime change” in European AD policy, a dummy variable was added equalling one for the years 1995-1998 and zero for the years before.

2. Specification Analysis

Due to the discrete and non-negative nature of the dependent variable NUMBERNEW, the normal linear regression model cannot constitute a valid data generating process. Rather, the formally correct way is to use a count data model, whose distributional assumptions account for the heteroscedastic and skewed distribution inherent to non-negative data and their discreteness. However, for large counts - like our dependent variable (see table 3) - the normal linear model might provide a reasonable approximation. Every regression equation was estimated[13] under the different distributional assumptions imposed by the Poisson, negative binomial maximum likelihood, and Poisson quasi-maximum likelihood (QML) count models.[14] As for the regressors, one proxy each was inluded for the domestic macroeconomic activity and the international influence via trade, as well as the growth rate of total industrial production in Japan. The lagged dependent variable, NUMBERNEW(t-1), was eliminated from a regression for redundancy reasons when insignificant. Combining each of the domestic and international variables gives nine regression equations per specification. This procedure serves three purposes. First, changing the control variables provides a sensitivity analysis of the regression results for the different regressors. Second, changing specification further analyses sensitivity, allowing for a comparison of the impact of each specification on the regression results. Third, specification tests enable one to draw inferences on the nature of the data generating process. Descriptive statistics of the variables are presented in table 4.

The regression results for the Poisson, the negative binomial, and the Poisson QML specification of each of the nine equations are given in tables A1, A2 and A3, respectively in

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13 The analysis was conducted using the computer package Econometric Views 3.1.

14 The Poisson regression model is the simplest count data model and can be considered as the benchmark model. However, if its assumptions are violated, estimation with this model cannot be efficient, and use of the Poisson standard errors would lead to biased inference. Therefore, we estimated each regression with these three specifications and conducted various tests in order to eliminate the appropriate specification for each regression. Details for why we used the specifications mentioned above, and the specification analysis itself are presented in Appendix A.

appendix A. It is noted here that a remarkable robustness of the regression results is found across the specifications with signs being identical in all cases but one, and the level of significance only differing for IPJAPGR(t-1) and NUMBERNEW(t-1).

Table 3: Dependent Variable Frequencies

Number of Investigations	Frequency	Relative Frequency	Cumulative Relative Frequency
20	1	5.56	5.56
21	1	5.56	11.11
24	1	5.56	16.67
25	1	5.56	22.22
27	1	5.56	27.78
29	1	5.56	33.33
33	1	5.56	38.89
36	1	5.56	44.44
38	1	5.56	50.00
39	2	11.11	61.11
40	1	5.56	66.67
43	2	11.11	77.78
45	1	5.56	83.33
48	1	5.56	88.89
49	1	5.56	94.44
58	1	5.56	100.00

Table 4: Descriptive Statistics of Variables

Variable	Mean	Std. Dev.	Min.	Max.
NUMBERNEW	35.89	10.59	20	58
GDPGR	2.13	1.19	-0.50	4.19
IPGR	1.67	2.45	-3.36	4.99
UER	0.24	0.79	-0.93	1.59
REER	0.90	0.08	0.79	1.10
TB	35.67	57.48	-45.10	131.80
IMPPEN	2.61	2.55	-2.73	6.19
IPJAPGR	2.04	4.27	-6.60	9.35

3. Regression Results

For ease of exposition, table 5 displays the regressions used for inference resulting from the specification analysis. As a general indicator of the goodness of fit of the model, the Wald or likelihood ratio tests[15] of the null hypothesis of joint insignificance of all included explanatory variables except the constant term strongly reject H_0 at the 1 % or 5 % (2.5 % in these cases) level of significance for all equations; i.e. the variation in the regressors explains to a significant degree the variation in the dependent variable.

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15 The likelihood ratio statistic is not valid for the Poisson QML model. Therefore, for the corresponding equations (1-4 in table 5) the Wald statistic was calculated using the computer package Gauss, for conducting a Wald test of the same null hypothesis. Recognizing that the coefficient vector is asymptotically normally distributed so that its square product with the inverse of the variance covariance matrix is $\aleph^2$-distributed, the Wald statistic can be calculated. While the LR and the Wald test statistic are asymptotically equivalent, they can lead to different results for small samples, where the Wald test has a higher probability of rejection under the null hypothesis. However, in our case the Wald statistics reject H_0 with a high significance level. (For a detailed discussion see *Berndt and Savin (1977)* and *Evans and Savin (1982)*).

Table 5: Dominating Specifications used for Inference

	CONSTANT	GDPGR(t-1)	IPGR(t-1)	UER(t-1)	REER(t-1)	TB(t-1)	IMPPEN(t-1)	IPJAPGR(t-1)	NUMBER-NEW(t-1)	LogLikelihood	LR, W[a]	df[b]	Iterations[c]
(1)	3.770303***	-0.193091***			0.115122			0.039152*			11.929297***	3	5
	(0.502751)	(0.061057)			(0.528063)			(0.020914)					
(2)	3.872307***	-0.186479***				-0.000238		0.036919			10.650348**	3	3
	(0.073671)	(0.071714)				(0.000783)		(0.024793)					
(3)			-0.100371***		-0.170957			0.040598			10.184560**	3	5
	(0.564083)		(0.032795)		(0.631012)			(0.022833)					
(4)			-0.100580***			7.52E-05		0.041381			9.6566531**	3	3
	(0.077765)		(0.037800)			(0.000735)		(0.026495)					
(5)	3.643815***			0.320528***	0.154134			0.041423***	-0.011402***	-63.25687	23.68043***	4	5
	(0.505102)			(0.068697)	(0.495513)			(0.013956)	(0.004070)				
(6)	3.795490***			0.319425***		-1.15E-05		0.041007***	-0.011692***	-63.30498	23.58422***	4	3
	(0.136863)			(0.078711)		(0.000715)		(0.015289)	(0.004003)				
(7)	3.807706***	-0.105191					-0.044178	0.034090**		-63.99732	22.19954***	3	3
	(0.086326)	(0.070650)					(0.028051)	(0.014841)					
(8)	3.671735***		-0.054780				-0.043790	0.035090**		-63.96534	22.26350***	3	3
	(0.055844)		(0.036291)				(0.028049)	(0.015160)					
(9)	3.825411***			0.214457*			-0.031417	0.037684***	-0.009403**	-62.84652	24.50113***	4	3
	(0.138702)			(0.129786)			(0.032825)	(0.014595)	(0.004608)				

Notes:
Equations (1)-(4) are taken from table A3 (Poisson QML), equations (5)-(9) are taken from table A1 (Poisson) in appendix A.
Standard errors (for Poisson) or Huber-White robust standard errors (for Poisson QML) in parentheses.

*, **, *** indicate significance at the 10 %, 5 % and 1 % level of significance respectively, using a two-tailed test.

[a] Likelihood ratio test statistic (for regressions 5-9) or Wald test statistic (for regressions 1-4) of H_0 : *joint insignificance of all regressors except the constant* against H_1 : *joint significance*. Asteriks indicate rejection of H_0. LR and W are asymptotically $\aleph^2$-distributed with q degrees of freedom, where q = number of restrictions.

[b] Degrees of freedom.

[c] Number of iterations completed for convergence.

Autocorrelation of the residuals was tested for up to the twelfth lag with the Ljung-Box-Q-statistic of H_0 : *residuals are serially uncorrelated*. In none of the regressions could H_0 be rejeced.

The regression results indicate that during the period 1980-98 the filing rate of antidumping and antisubsidy cases in the European Union was positively related to macroeconomic pressure. Looking at the individual variables, there is strong evidence for the hypothesis that pressures for antidumping protection in the EU are inversely related to the *domestic* macroeconomic situation. Equations 1 to 6 show high significance with the expected negative sign of the growth rate of real GDP (GDPGR(t-1)) and of total industrial production (IPGR(t-1)), and with the expected positive sign of the unemployment rate (percentage change from previous period, UER(t-1)). Thus, our results indicate that over the period under investigation, the more AD-investigations were initiated the lower was the growth of real GDP or of total industrial production, or the higher were the rates of unemployment. This suggests that in case of a macroeconomic downturn either industries file more petitions and/or the European CP-authority is more lenient in accepting requests for investigations.

Regarding *external* pressures approximated by the international trade position, we find strong insignificance of all three proxies. Thus our results do not indicate that a real (effective) appreciation of the EU countries‘ currencies implies a higher number of cases launched. Also, the insignificance of both the real effective exchange rate and the trade balance in particular suggest that the AD mechanism is not used for balance of payments reasons. Including import penetration (percentage change from previous period, IMPPEN(t-1)) leads to insignificance of the domestic macroeconomic variables, too, or, in the case of UER(t-1), strongly reduces the level of significance. This is, however, no evidence for a lack of robustness of the significance of the domestic macroeconomic variables but rather due to the high correlation of IMPPEN with GDPGR, IPGR and UER (0.81, 0.82, and –0.83, respectively), which inflates the standard errors of the collinear variables and thus reduces significance. Accordingly, likelihood ratio tests of the joint insignificance of IMPPEN(t-1) and each of the domestic macroeconomic proxies strongly rejected H_0 at the 1 % level of significance. The signs of the correlations suggest that at the aggregate level, import penetration might rather serve as an alternative proxy for macroeconomic activity. Comparing the impact of the domestic macroeconomic conditions and that of the international competitive factors on protectionist pressures, we can conclude that our evidence suggests a much more important impact of the former than of the latter.

Turning to the growth rate of industrial production in Japan, the results indicate quite a robust positive significance. Insignificance in equations 2 and 4 is likely due to multicollinearity introduced in the equations by IPJAPGR(t-1) and TB(t-1) (the correlation coefficient is 0.61). Deleting TB(t-1) from the regression leads to significance of IPJAPGR(t-1) at the 10 % level.[16] A likelihood ratio test of the null hypothesis of joint insignificance of the two variables in equations 2 and 4 highly rejected H_0 at the 5 % level of significance. The sign of IPJAPGR(t-1) is, however, contrary to what was expected. The positive sign might indicate that despite the inverse correlation of the growth rate of total industrial production and the volume of total exports, there may be a positive relation to the share of exports to the European Union. This could be the case if exports to the EU were mainly in goods with a low price elasticity of demand in Japan, so that a fall in IPJAPGR(t-1) would not necessarily result in a surge of these exports. However, this reasoning is contradicted by the positive sign of the correlation of IPJAPGR and TB.[17] As it stands, the sign is rather puzzling, and this may be an interesting issue for further research, also on the sectoral level.

The sign of the lagged number of newly initialized antidumping investigations, NUMBERNEW(t-1), suggests, when significant, existence of the ‚depletion effect‘ *Leidy (1997)* finds for the US. However, the significance is not robust to changes in the explanatory variables.

Finally, our regressions do not support the view that changes in the European trade defensive policy regulations agreed on in the Uruguay Round have exerted a significant change in the course of contingent protection policy in the EU as the coefficient of the included Uruguay Round dummy was insignificant[18] in all equations.[19]

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16 Deleting IPJAPGR(t-1), however, leaves TB(t-1) insignificant suggesting that the insignificance of TB(t-1) is not due to the correlation.

17 Also, this reasoning implies an effect of IPJAPGR(t-1) via the trade balance which, however, is found to be insignificant.

18 Due to their insignificance the regression results are not reported but available on request. The dummy variable was added to the regressions displayed in table 5.

19 Of course, however, tests for structural change in regressions with a low number of observations should be interpreted even more carefully than the regressions themselves, so it would be interesting to repeat this exercise in some years time when more observations will be available.

IV. Conclusions

The objective of this paper was to investigate the impact of the macroeconomic conditions on the pressure for contingent protection in the European Union. Similar to previous studies for the United States, we distinguished between domestic and external pressures. Our main results indicate that the domestic macroeconomic situation is strongly inversely related to pressures for contingent protection approximated by the number of newly initiated antidumping and antisubsidy cases. This result is robust to changes in the proxies for the macroeconomic pressure. However, with respect to external pressures all proxies were found to be insignificant. We also attempted to test for potential effects of the legislative changes implemented in 1995 after the Uruguay Round agreements. Our results suggest that (so far) there has been no significant change in the course of contingent protection policy in the EU as a result of these decisions. However, the investigation period 1980-98 only covers 18 annual observations so it may be worth repeating the exercise in a couple of years when more observations are available.

The regression results are remarkably robust to changes in the underlying model specification. Different count data model specifications were employed in order to account for the discrete and non-negative nature of the dependent variable, and various tests were conducted to derive the appropriate specification for each of the estimated equations.

Concluding, similar to evidence for the United States, the presented estimations indicate a strong impact of the domestic macroeconomic situation on the pressure for contingent protection in the European Union. External factors do, however, not seem to play a major role.

Appendix A: Specification Analysis

As outlined in section IV.2, the normal linear regression model cannot constitute a valid data generating process for discrete non-negative data. In the following, a specification analysis is conducted the result of which are the regressions used for inference displayed in table 5. The simplest count data model, the Poisson model, is the starting point of the analysis. Its probability density function is given by

$$\text{Prob}\left(Y_t = y_t\right) = \frac{\lambda_t^{y_t} e^{\lambda_t}}{y_t!} \;, \qquad t=1,\ldots,T \qquad \text{(A1)},$$

where λ_t denotes the Poisson parameter equal to the mean and the variance of the Poisson distribution. Typically, the Poisson regression model is given by

$$\ln \lambda_t = \boldsymbol{\beta}' \boldsymbol{x}_t \qquad \text{(A2)},$$

where $\boldsymbol{x}_t$ is the (1 x k) vector of regressors and β is the (k x 1) vector of coefficients. The parameters can be estimated with maximum likelihood techniques.

The Poisson maximum likelihood estimator resulting from maximization is consistent and efficient provided the conditional meanfunction is correctly specified and the conditional distribution of the dependent variable y_t is Poisson. If, however, the underlying distribution is not Poisson, the Poisson estimator, even though still consistent, will no longer be efficient, and use of the Poisson standard errors would lead to biased inference. The empirically most relevant case is a violation of the Poisson restriction that the mean must equal the variance. Most commonly, the violation of this assumption will be such that the data are characterized by overdispersion, i.e. the conditional variance exceeds the conditional mean.[20] In this case, the negative binomial model is an often used alternative to the Poisson model since its variance always exceeds the mean and it can so potentially accommodate for overdispersion. The negative binomial regression model takes the form[21]

$$\log \mu_t = \boldsymbol{\beta}' \boldsymbol{x}_t + \eta_t \qquad \text{(A3)},$$

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[20] The consequences of either over- or underdispersion (consistency but inefficiency, biased variance covariance matrix) resemble those of heteroscedasticity in the normal linear regression model.

where μ_t is the conditional mean, η_t reflects the specification error as in the normal linear regression model, with *exp(*η_t*)* Gamma distributed. If, however, the underlying distribution is not negative binomial either, the negative binomial maximum likelihood estimator will be both inefficient and inconsistent. Provided the mean is correctly specified Poisson quasi-maximum likelihood (QML) estimation will then yield consistent estimators even if the distribution is incorrectly specified.[22]

In order to test for the validity of the Poisson assumptions, three different tests were employed: those suggested by *Cameron and Trivedi (1990)* and by *Wooldridge (1996*) for testing mean variance equality and a likelihood ratio test of the Poisson against the negative binomial model.[23] The latter exploits the fact that the Poisson distribution is obtained as a parametric restriction of the negative binomial distribution.[24] Table 7 lists the results of the three tests. They indicate that the Wooldridge test is more restrictive than the test by Cameron and Trivedi, rejecting the Poisson model for equations 2, 3 and 4 at the 10 % level of significance, albeit very marginally for equation 2. The results are basically confirmed by the likelihood ratio test with the exception of equation 1 for which the test rejects the Poisson model in favour of the negative binomial model. However, the Wooldridge test does not reject the Poisson model only marginally here. For inferences, the Poisson specification was only accepted if neither of the tests were significant (equations 5-9). For reasons outlined above the Poisson QML estimation was used for the other equations (1-4).[25]

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21 For an in-depth formal treatment of both models see *Frome, Kutner and Beauchamp (1973), Hausman, Hall and Griliches (1984), Cameron and Trivedi (1986),* or *Winkelmann (1997).*

22 For a formal treatment of quasi-maximum likelihood estimation see *Gourioux, Monfort and Trognon (1984a, 1984b).*

23 The test by *Cameron and Trivedi (1990)* is based on an auxiliary ordinary least squares regression of the sum of squared residuals and the actual values on the square of the fitted values, $\left(e_0^2 - y\right)$ on $\hat{y}^2$. *Wooldridge (1996)* suggests to regress the standardized residuals minus one on the fitted values, $\left(e_s - 1\right)$ on $\hat{y}$. Rejection of the null with positive coefficient of the respective regressor indicates overdispersion in the data, i.e. the variance exceeds the mean.

24 For the model equations presented above, $\eta_t = 0$.

25 Even though the Poisson model is rejected in equations 1-4, it is interesting to note that it does quite well in this application in general as compared to microeconometric applications, given the relatively low level of significance of rejection (10%). This is basically due to the values of the dependent variable being relatively high, and a too large number of zeros playing no role.

Table A1: Specification Tests

	Cameron and Trivedi [a]	Wooldridge[b]	LR[c]
(1)	0.010258	0.018390	3.48856*
	(0.009010)	(0.010782)	
(2)	0.010427	0.018351*	3.46394*
	(0.008888)	(0.010490)	
(3)	0.013425	0.018820*	3.50468*
	(0.008538)	(0.009528)	
(4)	0.012911	0.018729*	3.50388*
	(0.009094)	(0.010114)	
(5)	0.019592	0.014386	2.22294
	(0.012706)	(0.010902)	
(6)	0.018956	0.014743	2.28930
	(0.012380)	(0.011123)	
(7)	0.005243	0.014108	2.34924
	(0.009658)	(0.011130)	
(8)	0.006438	0.014066	2.32510
	(0.009412)	(0.010443)	
(9)	0.013717	0.012721	1.78714
	(0.011700)	(0.010463)	

Notes:
Standard errors in parantheses.
*, **, *** indicate significance at the 10 %, 5 % and 1 % level of significance respectively, using a two-tailed test.
[a], [b] Cameron and Trivedi as well as Wooldridge test of the Poisson hypothesis of mean-variance-equality (H_0). For a description, see footnote x. Asteriks indicate rejection of H_0.
[c] Likelihood ratio test statistic of H_0 : *Poisson model* against H_1 : *negative binomial maximum likelihood model*.
Asteriks indicate rejection of H_0. LR is asymptotically $\aleph^2$-distributed with q degrees of freedom, where q = number of restrictions (here, q = 1 for all 9 tests).

Tables A1, A2 and A3 display the regression results of the Poisson, the negative binomial maximum likelihood, and the Poisson QML models, respectively.

Table A2: Poisson Regression Results

	CONSTANT	GDPGR(t-1)	IPGR(t-1)	UER(t-1)	REER(t-1)	TB(t-1)	IMPPEN(t-1)	IPJAPGR(t-1)	NUMBER-NEW(t-1)	LogLikelihood	LR[a]	df[b]	Iterations[c]
(1)	3.770303***	-0.193091***			0.115122			0.039152***		-65.21377	19.76664***	3	5
	(0.456987)	(0.043578)			(0.489892)			(0.014412)					
(2)	3.872307***	-0.186479***				-0.000238		0.036919**		-65.18104	19.83210***	3	3
	(0.075791)	(0.048053)				(0.000685)		(0.015389)					
(3)	3.782119***		-0.100371***		-0.170957			0.040598***		-65.12687	19.94043***	3	5
	(0.463989)		(0.022859)		(0.504273)			(0.014611)					
(4)	3.623983***		-0.100580***			7.25E-05		0.041381**		-65.17913	19.83591***	3	3
	(0.050390)		(0.025978)			(0.000724)		(0.016129)					
(5)	3.643815***			0.320528***	0.154134			0.041423***	-0.011402***	-63.25687	23.68043***	4	5
	(0.505102)			(0.068697)	(0.495513)			(0.013956)	(0.004070)				
(6)	3.795490***			0.319425***		-1.15E-05		0.041007***	-0.011692***	-63.30498	23.58422***	4	3
	(0.136863)			(0.078711)		(0.000715)		(0.015289)	(0.004003)				
(7)	3.807706***	-0.105191					-0.044178	0.034090**		-63.99732	22.19954***	3	3
	(0.086326)	(0.070650)					(0.028051)	(0.014841)					
(8)	3.671735***		-0.054780				-0.043790	0.035090**		-63.96534	22.26350***	3	3
	(0.055844)		(0.036291)				(0.028049)	(0.015160)					
(9)	3.825411***			0.214457*			-0.031417	0.037684***	-0.009403**	-62.84652	24.50113***	4	3
	(0.138702)			(0.129786)			(0.032825)	(0.014595)	(0.004608)				

Notes:
Standard errors in parentheses.
*, **, *** indicate significance at the 10 %, 5 % and 1 % level of significance respectively, using a two-tailed test.
[a] Likelihood ratio test statistic of H_0 : *joint insignificance of all regressors except the constant* against H_1 : *joint significance*. Asteriks indicate rejection of H_0. LR is asymptotically $\aleph^2$-distributed with q degrees of freedom, where q = number of restrictions.
[b] Degrees of freedom.
[c] Number of iterations completed for convergence.
Autocorrelation of the residuals was tested for up to the twelfth lag with the Ljung-Box-Q-statistic of H_0 : *residuals are serially uncorrelated*. In none of the regressions could H_0 be rejeced.

Table A3: Negative Binomial Maximum Likelihood Regression Results

	CONSTANT	GDPGR(t-1)	IPGR(t-1)	UER(t-1)	REER(t-1)	TB(t-1)	IMPPEN(t-1)	IPJAPGR(t-1)	NUMBERNEW(t-1)	Shape[a]	LogLikelihood	LR[b]	df[c]	Iterations[d]
(1)	3.816804***	-0.188143***			0.055063			0.038435**		-3.772419***	-63.46949	23.25520***	4	5
	(0.644068)	(0.059013)			(0.689839)			(0.019932)		(0.776256)				
(2)	3.864343***	-0.182436***				-0.000204		0.036651*		-3.778067***	-63.44907	23.29604***	4	6
	(0.106217)	(0.065027)				(0.000939)		(0.021346)		(0.778462)				
(3)	3.842349***		-0.097341***		-0.238245			0.039289**		-3.785601***	-63.37453	23.44512***	4	5
	(0.643616)		(0.030238)		(0.699814)			(0.019967)		(0.776654)				
(4)	3.622160***		-0.097162***			0.000103		0.040084*		-3.780182***	-63.42719	23.33981***	4	4
	(0.070004)		(0.034585)			(0.000990)		(0.022170)		(0.776855)				
(5)	3.695885***			0.320889***	0.115160			0.042509**	-0.011952**	-4.082637***	-62.14540	25.90337***	5	5
	(0.664165)			(0.087733)	(0.658910)			(0.018478)	(0.005223)	(0.911491)				
(6)	3.807141***			0.321757***		2.19E-05		0.042444**	-0.012164*	-4.071238***	-62.16033	25.87351***	5	4
	(0.177702)			(0.100865)		(0.000920)		(0.020144)	(0.005163)	(0.902066)				
(7)	3.803362***	-0.104638					-0.042388	0.033891*		-3.978616***	-62.82270	24.54878***	4	7
	(0.114798)	(0.091435)					(0.036287)	(0.019401)		(0.889982)				
(8)	3.667781***		-0.053778				-0.041774	0.034458*		-3.985834***	-62.80279	24.58859***	4	6
	(0.073811)		(0.046225)				(0.036332)	(0.019556)		(0.893852)				
(9)	3.834232***			0.228389			-0.026997	0.038900**	-0.010134*	-4.181558***	-61.95295	26.28828***	5	5
	(0.175863)			(0.164947)			(0.041294)	(0.018786)	(0.005857)	(0.986723)				

Notes:
Standard errors in parentheses.
*, **, *** indicate significance at the 10 %, 5 % and 1 % level of significance respectively, using a two-tailed test.
[a] Mixture parameter of the negative binomial model.
[b] Likelihood ratio test statistic of H_0 : *joint insignificance of all regressors except the constant* against H_1 : *joint significance*. Asteriks indicate rejection of H_0. LR is asymptotically $\aleph^2$-distributed with q degrees of freedom, where q = number of restrictions.
[c] Degrees of freedom.
[d] Number of iterations completed for convergence.
Autocorrelation of the residuals was tested for up to the twelfth lag with the Ljung-Box-Q-statistic of H_0 : *residuals are serially uncorrelated*. In none of the regressions could H_0 be rejeced.

Table A4: Quasi-Maximum Likelihood Regression Results

	CONSTANT	GDPGR(t-1)	IPGR(t-1)	UER(t-1)	REER(t-1)	TB(t-1)	IMPPEN(t-1)	IPJAPGR(t-1)	NUMBERNEW(t-1)	W[a]	df[b]	Iterations[c]
(1)	3.770303***	-0.193091***			0.115122			0.039152*		11.92930***	3	5
	(0.502751)	(0.061057)			(0.528063)			(0.020914)				
(2)	3.872307***	-0.186479***				-0.000238		0.036919		10.65034**	3	3
	(0.073671)	(0.071714)				(0.000783)		(0.024793)				
(3)	3.782119***		-0.100371***		-0.170957			0.040598		10.18456**	3	5
	(0.564083)		(0.032795)		0.631012			(0.022833)				
(4)	3.623983***		-0.100580***			7.52E-05		0.041381		9.656653**	3	3
	(0.077765)		(0.037800)			(0.000735)		(0.026495)				
(5)	3.643815***			0.320528***	0.154134			0.041423*	-0.011402***	30.68641***	4	5
	(0.621058)			0.093072	(0.566514)			(0.017342)	(0.004418)			
(6)	3.795490***			0.319425***		-1.15E-05		0.041007*	-0.011692***	22.94027***	4	3
	(0.132887)			(0.115522)		(0.000751)		(0.021574)	(0.003979)			
(7)	3.807706***	-0.105191					-0.044178	0.034090*		19.31654***	3	3
	(0.083269)	(0.090700)					(0.031230)	(0.017841)				
(8)	3.671735***		-0.054780				-0.043790	0.035090*		17.35031***	3	3
	(0.064812)		(0.047151)				(0.034433)	(0.018849)				
(9)	3.825411***			0.214457			-0.031417	0.037684**	-0.009403*	23.51310***	4	3
	(0.123996)			(0.167753)			(0.040993)	(0.015635)	(0.005179)			

Notes:
Huber-White robust standard errors in parentheses.

*, **, *** indicate significance at the 10 %, 5 % and 1 % level of significance respectively, using a two-tailed test.

[a] Wald test statistic of H_0 : *joint insignificance of all regressors except the constant* against H_1 : *joint significance*. Asteriks indicate rejection of H_0. W is asymptotically $\aleph^2$-distributed with q degrees of freedom, where q = number of restrictions.

[b] Degrees of freedom = number of restrictions.

[c] Number of iterations completed for convergence.

Autocorrelation of the residuals was tested for up to the twelfth lag with the Ljung-Box-Q-statistic of H_0 : *residuals are serially uncorrelated*. In none of the regressions could H_0 be rejected.

Appendix B: Data

Data Sources

Commission of the European Community, Annual Reports of the Commission of the European Communities on the Community's Antidumping and Antisubsidy Activities (1983-1998): data on the number of newly initiated antidumping and antisubsidy cases.

OECD Main Economic Indicators database: data on real GDP, imports of goods and services (both in 1990 US-$), total industrial production for the EU and Japan, the real effective exchange rate, and the consumer price index (all items) (all index numbers, 1995=100).

OECD Economic Outlook (June 1998 and 1999): data on the trade balance (in US-$, for the analysis deflated to 1990 constant prices).

Sachverständigenrat zur Begutachtung der gesamtwirschaftlichen Entwicklung (Council of Economic Experts) (1999): data on total numbers of unemployed and employed persons.

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