Determining Military Expenditures: Arms Races and Spill-Over Effects in Cross-Section and Panel Data

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Abstract

This paper considers the determinants of military spending, building on an emerging literature that estimates military expenditure demand functions in cross-section and panel data, incorporating 'arms-race' type effects. It updates Dunne and Perlo-Freeman (2003b) using the SIPRI military expenditure database for the period 1988-2003, finding broadly similar results. It also shows differences in results across panel methods, particularly the within and between estimates and illustrates the importance of recognising and modelling dynamic processes within panel data. Heterogeneity is also found to be an important issue and when countries are broken up into groups on the basis of per capita income there is no obvious systematic pattern in the results. This is seen to imply that the demand for military spending, even between two mutually hostile powers, may depend on the whole nature of the relationship between them (and other countries and events in the region), and not simply Richardsonian action-reaction patterns.

Keywords: Military Spending; Demand; Arms races; Spillovers; Panel data

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1.   Introduction

Military spending can impose considerable economic burdens on countries, particularly developing ones, making it important to understand what factors are determining it, whether strategic or economic. A particular concern is the degree to which particular countries' are influencing neighbours and producing local arms races. A growing literature attempts to identify the factors that influence the evolution of military spending and conflict across countries. It has moved beyond the earlier cross sectional analyses to use panel data methods, which incorporate both time series and cross sectional effects, a recent example of which is Collier & Hoeffler (2007). Initially this was simply a move to pooling data or using static panel methods, but the development and understanding of methods to analyse dynamic panels has led to their increasing use, as in Dunne & Perlo-Freeman (2003b).

This paper continues this line of investigation. It considers the issues raised in Dunne and Smith (2007) on how the military spending of other countries can impact upon a country's military spending, the issues involved in using panel data methods and the importance of heterogeneity.. The next section presents the basic arms race, action-reaction theoretical model and reviews the econometric issues relating to cross-section and panel data. Section 3 then develops the work of (Dunne, Perlo-Freeman and Smith, 2008), by using SIPRI military expenditure data for the period 1988-2003 to update the model of the demand for military spending in Dunne and Perlo-Freeman (2003b) and comparing the results using the different available econometric techniques. The degree and nature of heterogeneity is then considered in Section 4, with Section 5 proving some conclusions.

2. Arms Race Models and Strategic Effects

The starting point for any analysis of the impact of one countries arms procurement or military spending on anothers is the basic Richardson (1960) 'arms race' model. This supposes two countries whose military expenditure/level of arms/military capability, m_1 and m_2, are related at time t by the equations:

$$\frac{dm_1(t)}{dt} = a_1 + b_1 m_2(t) - c_1 m_1(t)$$

$$\frac{dm_2(t)}{dt} = a_2 + b_2 m_1(t) - c_2 m_2(t)$$

Where a_i are exogenous 'grievance' terms, b_i are 'reaction' terms, whereby each country responds to the military capability of the other, and c_i are 'fatigue' terms, usually representing some internal limitations on a country's military spending/capability. Alternatively, a discrete formulation can be made using difference equation,

$$\Delta m_{1t} = a_1 + b_1 \ m_{2t} + c_1 m_{1,t-1}$$

$$\Delta m_{2t} = a_2 + b_2 \ m_{1t} + c_2 m_{2,t-1}$$

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This basic model has been developed theoretically and empirically in a variety of ways[1] The search for clear empirical evidence of 'arms races' has, however, met with rather limited success, with even apparently obvious examples as the Cold War superpower arms race proving ambiguous and very much dependent on specification[2].

A number of authors, including Dunne and Perlo-Freeman (2003a and 2003b) and Collier and Hoeffler (2004), have sought to generalise the concept of an arms race by looking at the demand for military expenditure across a large group of countries, using either cross-section or panel data, incorporating a range of economic, political and security variables, and including variables for the aggregate military expenditure of neighbours and rivals. In addition, Murdoch and Sandler (2002, 2004) have focussed on the 'spillover effects' of military spending.

To illustrate the issues involved consider a cross-section of countries $i = 1, 2, .., N$, with the average value for some military measure m_i over some time-period and other determinants

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[1] Including the explicit modelling of rational economic decision-making, different dynamic specifications, game theory approaches, and empirically with the use of approaches such as co-integration (Dunne and Smith, 2007).
[2] India and Pakistan provides one of the few examples where researchers have been able to provide consistent evidence of a Richardsonian arms race. Even then, Oren (1994) has offered an alternative approach, based on hostility levels between the two countries, under which the apparent arms race disappears. Numerous attempts have been made to estimate arms races for Turkey and Greece, using a variety of theoretical and econometric models, without clear evidence of an arms race emerging. (E.g. Dunne, Nikolaidou and Smith (2003), and Smith, Sola and Spagnolo (2000).)

x_i. If one particular country i feels threatened by an alliance of j and k; the equations for these 3 (out of N) observations take the form:

$$m_i = \alpha + \beta^e (m_j + m_k) + \gamma' x_i + \varepsilon_i$$

$$m_j = \alpha + \beta^e m_i + \beta^a m_k + \gamma' x_j + \varepsilon_j$$

$$m_k = \alpha + \beta^e m_i + \beta^a m_j + \gamma' x_k + \varepsilon_k$$

where β^e measures the arms race effect from an enemy and β^a the spillover effect from an ally. Given data on N countries and knowledge of threats and alliances, this model could be estimated by say OLS.

There are obvious practical problems; how to determine the strategic effects, the pattern of threats and alliances, how to aggregate if adding allies expenditure is not appropriate, what time period to average over, and which military measure to use,. An early attempt to deal practically with such strategic effects was provided by Rosh (1988), who introduced the concept of a security web. Instead of dyadic relationships, neighbours and other countries (such as regional powers) that can affect a nation's security were considered part of a country's Security Web and their military spending was aggregated[3]. More recently, Dunne and Perlo-Freeman (2003a) developed this approach as discussed below.

There are also econometric issues[4]. Collier and Hoeffler (2004) estimate arms race multipliers by considering the two country case:

$$m_i = a_i + bm_j$$

An exogenous unit increase in a_i would cause an increase in the other countries military expenditure by bm_j feeding back on m_i and giving a total effect of $1/(1-b)$ on spending by

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[3] Rosh calculates the degree of militarisation of a nation's Security Web by averaging the military burdens of those countries in the web, finding it to have a significant positive effect on a country's military burden.

[4] There is an issue of simultaneity, but it is not the standard one, where the right hand side variable may be determined by another equation. All N military expenditures are determined by the same equation. It is more like a lagged dependent variable issue and the conditions for consistent estimation by OLS are the same as that case:

country i. They distinguish own expenditure and neighbour expenditure arms race multipliers.

Using cross-section data, rather than time-series for individual countries increases the sample size considerably and allows one to measure the effect of variables which tend not to vary very much within countries. The cost of this benefit is the assumption that the arms race and alliance parameters are the same across all countries. This assumption can be weakened and allowing the coefficients to differ across countries and collect the independent variables in a vector z_i gives the model

$$m_i = \alpha_i + \beta_i' z_i + \varepsilon_i$$

with $\alpha_i = \alpha + v_i$, $\beta_i = \beta + \eta_i$. When $\alpha = E(\alpha_i)$, $\beta = E(\beta_i)$ and η_i and v_i are random, independent of the regressors, OLS will provide consistent estimates of β. These assumptions are, however, quite strong and with the availability of panel data models of the form

$$m_{it} = \alpha_i + \beta_i' z_{it} + \varepsilon_{it}$$

can be estimated and the homogeneity or independence assumptions tested (Dunne and Smith, 2007). With panel data one can employ a larger sample and allow for heterogeneity in the responses of different countries.

Concentrating on a single equation[5], we have data on a scalar dependent variable m_{it}, military expenditure, for a sample of countries $i = 1, 2, ..., N$ and years $t = 1, 2, ..., T$ and data on a vector of k exogenous variables x_{it}, which does not include unity for an intercept, but which includes other countries military expenditure. Suppose that we are interested in measuring the linear effect of x_{it} on m_{it}. This is a static model, in that x_{it} does not include lagged dependent variables; we discuss dynamics below.

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m_i can be treated as predetermined in the equation for m_j if $E(\varepsilon_i \varepsilon_i) = 0$. If there are correlated shocks to different countries or spatial serial correlation, this assumption will not hold (Dunne and Smith, 2007).

A central issue in the choice of estimator is the relative size of N and T. The traditional panel literature deals with cases where N is large and T small, maybe only two or three time periods. Asymptotic analysis is done letting $N \rightarrow \infty$. The time-series literature deals with the case where T is large and N small and asymptotics let $T \rightarrow \infty$. Recently there has been interest in panel time-series where N and T are of the same orders of magnitude and asymptotics let both $N \rightarrow \infty$ and $T \rightarrow \infty$ in some way. What estimators are appropriate in the three cases differs. Define the country and overall means as

$$\overline{m}_i = T^{-1} \sum_{t=1}^{T} m_{it}; \quad \overline{m} = (NT)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} m_{it}.$$

The total variation in the dependent variable is the sum of the within country variation and the between country variation

$$\sum_i \sum_t (m_{it} - \overline{m})^2 = \sum_i \sum_t (m_{it} - \overline{m}_i)^2 + T \sum_i (\overline{m}_i - \overline{m})^2,$$

similarly for the regressors (Dunne and Smith, 2007).

The main panel estimators differ in how they treat the within and between variation. The estimators include:

- Pooled OLS which gives the within and between variation equal weight, and uses least squares on:

  $$m_{it} = \alpha + \beta' x_{it} + u_{it};$$

  where β is a $k \times 1$ vector of slope coefficients;

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[5] To focus on this, we will ignore feedback issues, to concentrate on a single equation. When one allows for panel VARS and VECMs, the situation becomes more complicated.

- The within estimator (also known as the one way fixed effects, least squares dummy variables and a variety of other names) which allows intercepts to differ across countries but constrains the slopes to be the same

$$m_{it} = \alpha_i + \beta' x_{it} + u_{it}$$

    this only uses the within variation and is equivalent to OLS on

$$(y_{it} - \bar{y}_i) = \beta'(x_{it} - \bar{x}_i) + u_{it}$$
$$\text{since } \alpha_i = \bar{y}_i - \beta' \bar{x}_i;$$

- The between or cross-section estimator, which only uses the cross-section variation in the country means:

$$\bar{m}_i = \alpha + \beta' \bar{x}_i + \bar{u}_i;$$

- The two way fixed effects estimator which constrains slopes to be the same, but allows intercepts to vary freely both over country and year:

$$m_{it} = \alpha_i + \alpha_t + \beta' x_{it} + u_{it}$$

    this allows for a completely flexible trend, or unobserved common factor, which impacts on each country by the same amount;

- The random coefficient model (RCM), , which allows all the parameters to differ over countries:

$$m_{it} = \alpha_i + \beta_i' x_{it} + u_{it}$$

and calculates weighted averages of the individual time-series estimates $\hat{\beta}_i$ the weights, W_i being based on the variances of the $\hat{\beta}_i$ (Swamy, 1970).

The random effects estimator assumes that slopes are identical and intercepts are randomly distributed independently of the regressors. It calculates the optimal combination of within and between variation under these assumptions[6].

Strictly speaking, we should distinguish the parameters: the β in the within equation is a different parameter from the β in the between equation. The distinction is important because the cross-section may be measuring something quite different from the time-series[7]. In static models, if the coefficients, α_i and β_i are randomly distributed, independently of the x_{it}, all these estimators will produce unbiased estimators of the expected values of the coefficients $E(\alpha_i)$ and $E(\beta_i)$. However, the independence assumption may not hold and the cross-section (between country) effects can be very different from the time-series (within country) effects. The between estimate β_B and the within estimate β_W will differ if the α_i are correlated with the $\overline{x}_i$ and the two estimates may even be of opposite sign

A further issue arises with dynamic models since the within (fixed effect) estimator of

$$m_{it} = \alpha_i + \beta' x_{it} + \lambda m_{it-1} + u_{it}$$

is consistent for large T, but is not consistent for fixed T, large N. In this case the coefficient of the lagged dependent variable is biased downwards. This is the standard small T bias of the OLS estimator in models with lagged dependent variables. There are a variety of instrumental variable estimators for this case. However, if the true model is heterogeneous

$$m_{it} = \alpha + \beta_i' x_{it} + \lambda_i m_{it-1} + u_{it}$$

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[6] The TSCS (time-series cross-section) estimator allows for between group heteroskedasticity and within group serial correlation. This estimator is rarely used in economics, because it treats the symptoms of misspecification (heteroskedasticity and serial correlation) rather than their likely causes (parameter heterogeneity and dynamic misspecification), though it is popular elsewhere (Dunne and Smith, 2007).

[7] Dunne and Smith (2007) also discuss a number of other estimators. The Zellner SURE estimator allows for non-zero between group covariances and possibly heterogeneous slopes. This is only feasible when T>N, and only appropriate when the factors causing between group dependence are not correlated with the regressors. There are other estimators which can deal with between group dependence when N>T, see the review in Smith and Fuertes (2005).

and homogeneity of the slopes is incorrectly imposed, the within estimator is not consistent even for large T. The coefficient of the lagged dependent variable is biased upwards towards unity (assuming the regressors are positively serially correlated as is usually the case). The RCM estimator is however consistent for large T, though it suffers the small T lagged dependent variable bias. Comparison of the various estimators, which are subject to different biases, can allow us to infer which biases are most important.

In the individual regressions,

$$m_{it} = \alpha_i + \beta_i' x_{it} + u_{it},$$

if the variables are integrated of order one, I(1) and are cointegrated (the error term u_{it} is I(0)) then the least squares estimate of $\hat{\beta}_i$ gives a super-consistent estimate of the long-run effect for large T. However, as noted above if the variables are I(1) but are not cointegrated (the error term is also I(1)), then $\hat{\beta}_i$ does not converge to β_i, as $T \to \infty$, but to a random variable. The regression is spurious. However, pooling or averaging over groups, can allow one to obtain a consistent estimate of an average long run effect from the levels regression. Thus the pooled or average estimates from static levels regressions may be of interest even if individual countries equations differ and do not cointegrate (Dunne and Smith, 2007).

Clearly the use of panel data methods can provide benefits to the researcher, but care is required to understand the nature of the models and how relevant they are for the particular study.

3. Empirical Analysis

To investigate these issues empirically we start from the model used in Dunne and Perlo-Freeman (2003a,b), and update the analysis, which means moving from using ACDA data on military expenditure to the SIPRI data, which is the largest consistent database available, c Covering 136 developing and developed countries from 1988-2003[89]. The results for the fixed

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[8] The SIPRI database goes up to 2005, but data for the Democracy variable used in the regressions is only available up to 2003.

effects regressions are presented in Table 1, with the dependent variable the log of military burden and the independent variables (with variable names in brackets):

- log of GDP (lgdp),
- log of population (lpop)
- War index variable[10] (war)
- Democracy variable using the Polity 4 dataset (polity)
- log of aggregate 'Security Web' military spending[11] (lsw)
- log of aggregate Rivals' military spending (a subset of the Security Web consisting of those powers considered hostile to the country in question)[12] (lriv)
- 'Great Power Enemy' dummy for countries that are considered to be an enemy of a 'great power' (US or USSR) (gpe)

In addition, a number of dummy variables are used:

- Cold War dummy covers the years 1988-90, to account for the possibility that the end of the Cold War led to a general reduction in threat perceptions. (cold)
- Cold Warrior dummy is set to 1 for the years 1988-90 for NATO and Warsaw Pact members only. (coldw)
- Iraq Crisis dummy is set to 1 for the years 1990-91 and 2002-03, to capture the effect of the two Iraq wars as 'regional bads' that may have increased threat perceptions, even for those countries not directly involved.[13] (iraq)

As a 'security web' is likely to have a different meaning for countries involved in the NATO and former Warsaw Pact alliance systems than for those who were not, two versions of the

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[9] Many studies (including those by Dunne and Perlo-Freeman) involve chaining together military spending data from different editions of the SIPRI and ACDA yearbooks, which is subject to inaccuracies due to periodic re-estimation of the data by the two organisations. The SIPRI military expenditure project has now however ensured that it has consistent data series for all countries in its database going back to 1988.

[10] An index running from 0 to 4, based on the Uppsala Department of Peace and Conflict Research database. The previous study had separate variables for External War and Civil War, but the effects of these were fairly similar, so the two types of war were combined in this study.

[11] A country's Security Web (based on the idea of Rosh (1988) is the set of counties that are considered to affect a given country's security. These are mostly neighbours, but may also include regional powers and sometimes more distant rivals (e.g. Cuba-South Africa during the Angola war). The military spending of superpowers is usually excluded from a country's Security Web, where it is considered that it would be impossible to defend against the superpower with conventional military expenditure.

[12] The classification of countries as mutually hostile is based on an analysis of a number of datasets, including the HIIK Database of Violent and Non-violent Conflicts, and the Dyamid datset of dyadic armed conflict events. The previous study also used a further subset, Enemies, but this latter variable was insignificant, suggesting that the effect of those classified as 'Enemies' and those merely considered 'Potential Enemies' was not significantly different. Thus only the broader class of Potential Enemies (or Rivals) is used here.

[13] These countries are: Egypt, Iran, Israel, Jordan, Kuwait, Lebanon, Oman, Saudi Arabia, Syria, Turkey, UAE, Yemen. (Iraq is not in the sample, as there is no data for Iraq in the SIPRI Military Expenditure Database).

model were estimated. One, using the full sample and excluding the Security Web variable, and the other including this variable but excluding most NATO and WP countries and of course the Cold Warrior dummy[14]. The results, presented in Table 1, are similar to Dunne and Perlo-Freeman (2003b), with the income elasticity of demand slightly less than unity, population negative and highly significant[15], war positive and highly significant, and democracy negative. However, the Rivals' military spending variable[16], while positive, is insignificant and small in absolute value, while the Security Web variable in Model 2 is both positive and highly significant[17]. Surprisingly, there is a strongly negative coefficient on the Great Power Enemy dummy, which may simply be an anomaly and the other results did not change significantly when this variable was removed[18].

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[14] Greece and Turkey are included, as it was considered that their Security Webs could be defined in terms of the Mid-East/Balkans systems. The former Soviet republics of the South Caucasus and Central Asia are also included.

[15] One interpretation of this result within a rational maximising model is that higher population creates additional civil needs faster than it creates additional security needs.

[16] Corresponding to the Potential Enemies variable in the previous study

[17] This may suggest that there are stronger 'emulation' effects than in the past, with countries seeking to keep up with their neighbours' purchases of the latest equipment aside from any threat, but another explanation is that, in the mostly post-Cold War environment where interstate war is less prevalent and less of a threat for most countries, threat perceptions may be based more on common regional threats, including terrorism, spillovers from conflicts in the region, etc. Examples of such common threats are seen in the Iraq crisis, Cold War and Cold Warrior dummies, which are all positive and significant, with quite large coefficients on the Iraq Crisis and Cold Warrior variables in particular.

[18] There are in fact very few countries in the sample whose GPE status has *changed* over the period Specifically, Angola, Cambodia, Turkey (all ceasing to be GPEs after 1990) and Serbia (after 2000). The characterisation of the first two is in any case quite marginal, and in the case of Turkey the ending of hostility with the Soviet Union is already captured by the end of the Cold War and it is only these cases that can be picked up by the fixed effects estimator. Thus the coefficient may simply be picking up idiosyncratic conditions in these countries.

Table 1: Fixed Effects Regression results 1988-2003

Dependent variable: log military burden.	Model 1 (all countries)		Model 2 (exc. NATO, WP other than Greece, Turkey)	
Variable	Coefficient	T-ratio	Coefficient	T-ratio
Log GDP (lgdp)	-0.052	-2.3**	-0.054	-2.1**
Log Population (lpop)	-0.45	-5.1***	-0.52	-5.1***
War	0.13	10.1***	0.14	9.7***
Democracy (polity)	-0.0078	-3.3***	-0.0050	-1.8*
Log Security Web military spending (lsw)			0.12	4.2***
Log Rivals' military spending (lriv)	0.0076	1.6	0.0062	1.2
Iraq crisis dummy (Iraq)	0.15	3.2***	0.13	2.6***
Cold War dummy (cold)	0.072	3.0***	0.059	2.2**
Cold Warrior dummy (coldw)	0.19	4.2***		
Great Power Enemy (gpe)	-0.69	-5.7***	-0.55	-4.3***
R^2 (within)	0.17		0.17	
No. observations	1856		1364	
No. countries	136		103	

Note: *** sig at 1%; ** sig at 5%

These results are from a fixed effects static model and there are three specific issues have to consider as discussed in section 2, the choice of panel estimation method, the treatment of dynamics and the degree of heterogeneity within the panel. The results in Table 2 illustrate the variation across the methods, with clear differences between the fixed effects, random effects and the OLS estimates. It is also clear that the simple pooled OLS provides somewhat misleading results. As expected there are also marked differences, between within and between estimators. As discussed these are likely to be measuring different things, but it is

interesting that the cross section finds no role for economic variables and that this consistent with other studies. There are also differences between static and dynamic results, with the lagged dependent variable significant. There is, however, an issue of using the fixed effects model with a lagged dependent variable with a limited number of years of data. The results do, however, serve to illustrate the likely importance of the dynamics.

Table 2: Summary results for Model with LSW

Results for models with lsw

Dependent variable : logburden

Variables	Static		Between		Dynamic		Long run	
	OLS	FE	RE		OLS	FE	OLS	FE
Lgdp	0.00	**-0.05**	-0.03	-0.01	0.00	**0.04**	0.01	**0.11**
Lpop	**-0.11**	**-0.52**	**-0.07**	-0.06	-0.01	**-0.17**	-0.09	**-0.41**
War	**0.16**	**0.14**	**0.14**	**0.17**	**0.03**	**0.08**	**0.38**	**0.20**
Polity	**-0.03**	-0.01	**-0.01**	**-0.03**	**0.00**	0.00	**-0.03**	-0.01
Lriv	**0.03**	0.01	**0.01**	0.03	0.00	0.00	-0.01	0.00
Iraq	**0.75**	**0.13**	**0.14**	**3.67**	**0.15**	**0.13**	**1.76**	**0.33**
Cold	0.05	**0.06**	**0.11**	-0.26	0.03	**0.07**	0.38	**0.17**
Gpe	-0.14	**-0.55**	**-0.36**	-0.21	0.01	-0.13	0.06	-0.32
Lsw	**0.13**	**0.12**	**0.12**	**0.09**	**0.01**	0.02	**0.13**	0.05
logburden1					**0.91**	**0.60**		
_cons	**1.32**	**9.60**	**1.73**	1.25	0.07	1.87	0.79	4.63
Rsq	0.28	0.17	0.15	0.03	0.92			
No. countries	103.00				101.00			
No. obs	1364.00				1251.00			

As has been noted, the Fixed Effects estimator does not allow for parameter heterogeneity in the slope coefficients, only in the intercepts. A random coefficients (mean group) estimator would deal with this problem, but we do not have a sufficient number of time periods for this to be feasible, and too many of the variables are time-invariant for too many countries. Where the variables are I(1) (as variables such as GDP, military spending and population typically are, at least over short periods) the FE estimator may nonetheless give consistent estimates of the average coefficients. (Dunne and Smith, 2007). The question remains of exactly how important heterogeneity and this is considered in the next section.

4. Heterogeneity

To give some idea of the importance of heterogeneity we divided the sample into four groups based on average GDP/capita over the sample, running the regressions above on each. For the bottom two groups we used model 2, as the security web is defined for most of these countries. For the richest group, we used model 1, as a high proportion of these countries are members of the alliance systems for which security web is not defined, and for the second highest group we used both. The results in Table 3 do indeed show that there is considerable heterogeneity in the determinants of milex. The poorest group showing the clearest effect of security web and rivals' milex, with positive and highly significant coefficients on both, and the coefficient on security web being as large as 0.37. There was a strong positive effect of war, and a negative effect of democracy as in the main sample, and a clear negative effect of GDP (on military burden). In contrast to the overall regression, population was clearly insignificant, though still with a negative sign. The within R^2 was quite high at 0.27.

Table 3 – Fixed Effects regression results for different income groups.

Variable	Low income	Mid-low	Mid-high (model 1)	Mid-high (model 2)	High income
Log GDP	-0.43*** (0.140)	-0.400*** (0.0913)	0.0522** (0.0259)	0.0287 (0.0266)	-0.981*** (0.0684)
Log pop.	-0.115 (0.248)	-0.0562 (0.203)	-0.858*** (0.239)	-0.824*** (0.259)	0.338** (0.159)
War	0.111*** (0.027)	0.0299 (0.0248)	0.156*** (0.0256)	0.157*** (0.0256)	0.0332 (0.0252)
Democracy	-0.0102** (0.00468)	0.00090 (0.00392)	-0.0273*** (0.00616)	-0.0229*** (0.00699)	0.00966 (0.0217)
Log Security Web	0.369*** (0.637)	-0.0226 (0.0395)	----------	0.163*** (0.0588)	------------
Log Rivals	0.0463*** (0.0126)	-0.00225 (0.00784)	-0.0206* (0.0107)	-0.0251** (0.0111)	-0.00004 (0.0064)
Cold war dummy	-0.0168 (0.0583)	0.0300 (0.0442)	0.0363 (0.0539)	0.0200 (0.0570)	-0.0453 (0.0294)
Cold Warrior dummy	-----	-------	-0.0735 (0.111)	--------	0.144*** (0.0384)
Iraq dummy	-0.243 (0.279)	0.0221 (0.100)	0.152* (0.0891)	0.123 (0.0894)	0.185*** (0.0420)
R^2 (within)	0.266	0.131	0.207	0.240	0.521
No. countries	34	31	33	24	33
No. obs.	420	395	447	329	517

Standard errors in brackets. * = sig. at 10% level ** = 5%, *** = 1%.

The second poorest group showed the least clear results, with only GDP significant and negative, all other variables being jointly insignificant. The coefficients on these variables were also of much smaller magnitude than for other groups, so that this is not just a matter of high standard errors due perhaps to lack of variation. This might suggest that there are further major heterogeneities within this group that are largely cancelling each other out to give no clear overall pattern. The R^2 was 0.13

The second richest group showed a mixed pattern; in model 2 there was a large (-0.88), negative and significant coefficient on population, again positive and significant coefficients on war and security web and negative and significant on democracy, but a *negative* coefficient on rivals' milex.[19] There are a few countries for which the Iraq dummy is meaningful (Iran, Turkey, Syria), and the coefficient is positive but insignificant. Model 1 included a larger number of countries (many of the former Warsaw Pact), and had similar results, except that GDP was now significant and slightly positive, and the Iraq dummy was also marginally significant. The R^2 was 0.20 for model 1, and 0.24 for model 2.

The richest group, which included mostly the developed world countries, Israel and the Gulf states and some of the richer Asian countries, showed a decidedly different pattern: the coefficient on GDP was highly significant and negative, in fact not significantly different from -1, implying that for these countries military spending is independent of income. Population was significant and positive, in contrast to the overall results. The only other significant variables were the 'Cold Warrior' and Iraq dummies. The R^2 was higher than the rest at 0.52. For these countries, then, military spending seemed primarily to relate to major security issues such as the Cold War and the Iraq crises. This in turn immediately implies heterogeneities, in that these dummies were not relevant for all countries in the group. This 'average' sub-panel result tells us very little about the determinants of military spending in (for example) Brunei, Cyprus, Finland, Singapore, South Korea or Sweden. Curiously, the Cold War dummy is not significant in any subgroup although it is significant in the overall results.

There no obvious a pattern to the differences between these groups, and indeed these results do not suggest that differences in income are the basis of heterogeneities. A regional sub-classification sheds no more light, other than obvious points such that it was in the Middle East that the Iraq crises were most relevant, and in Nato and the Warsaw Pact the Cold War. In other words, it would seem that heterogeneities must be looked for at the individual country level. This is not to say that the overall panel results are meaningless – they do suggest general tendencies that must apply in a majority of countries for which the issue is

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[19] As rivals are also included in the security web, this does not imply an overall negative result, as the positive coefficient on security web was much higher. However it does suggest that these countries respond less to the milex of rivals than to that of other neighbours.

pertinent; for example the general negative influence of democracy on military spending, one of the most consistent results in analysis of milex demand, but one which cannot be detected in this study for countries that have not changed their democratic status over the period; or perhaps the tendency of nearby countries to respond to common threat perceptions which may be suggested both by the Iraq dummy and the Security Web variable; but that which issues are pertinent at a given time, and in what way, will vary considerably from country to country.

To explore more specifically the arms race concept (proxied by the Rivals' milex effect) it is useful to get an idea of the different nature of the reactions by different countries to change in their rivals' military spending. Dunne & Perlo-Freeman (2003b) found this effect to be highly significant and robust in the whole sample, while this study has found a positive but insignificant result, but it may be that this masks very different patterns of response in different countries.

To investigate this, country-specific variables have been constructed, equal to the Rivals variable times the dummy for each country in question, for those countries who at some stage had non-zero values of this variable.[20] The other independent variables are kept as single variables for the whole sample. Thus, each country is allowed to have a different response to the aggregate military spending of its rivals, although homogeneity is still assumed for the other variables. A like response is also assumed for each country to its various different rivals, where it has more than one.

Again other variables give similar results to those in table 1, although population becomes insignificant, and some other variables become less significant. Out of 35 country-specific Rivals' military spending variables, 8 are positive and significant, 5 are negative and significant, 12 are positive and insignificant, 10 are negative and insignificant. (See results in Table 3 and figure 1). Coefficient values range from -3 to +2, although clearly these must be treated with extreme caution given the small T. Thus, in keeping with the lack of significance of the Rivals' military spending variable overall, there appears to be no consistent pattern to the influence of the variable on individual countries. On the other hand, a fair number of countries do appear to have positive and significant responses to rivals' milex, and in most of

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[20] The log variables were constructed as log(aggregate military spending + 1), so that 0 values are indeed transformed to 0.

these cases the rival(s) in question mostly have at least a positive (albeit insignificant) coefficient.

It should be noted that this was the model with the Security Web military spending variable; so there seems to be a general influence of neighbours' military spending, but no consistency with regard to there being an incremental influence from those countries who are considered rivals. Again, this may suggest emulation effects or common threat perceptions, but not necessarily arms races in the conventional sense. A similar exercise was conducted for the ACDA dataset used in Dunne & Perlo-Freeman (2003b) and this also showed a highly heterogenous pattern, but a much greater preponderance of positive coefficients on the country-specific Potential Enemies variables, in keeping with the significant positive coefficient on the general Potential Enemies variable (Dunne, Perlo-Freeman and Smith, 2008).

Table 3: Results for Country-specific Variables

Country	Potential Enemies Military Spending
Algeria	-ve ins.
Angola	-ve sig.
Argentina	-ve ins.
Armenia	-ve ins.
Bangladesh	-ve ins.
Bolivia	+ve ins.
Cameroon	+ve ins.
Chile	-ve ins.
China	-ve ins.
Cyprus	-ve sig.
Egypt	+ve ins.
Eritrea	+ve ins.
Ethiopia	**+ve sig.**
Greece	+ve ins.
India	+ve ins.
Iran	-ve sig.
Israel	-ve ins.
Jordan	-ve ins.
South Korea	+ve ins.
Mauritania	**+ve sig.**
Morrocco	+ve ins.
Mozambique	**+ve sig.**
Myanmar	+ve ins.
Nigeria	**+ve sig.**
Pakistan	-ve sig.
Peru	-ve ins.
Rwanda	+ve ins.
Saudi	+ve ins.

Senegal	+ve ins.
South Africa	**+ve sig.**
Sudan	**+ve sig.**
Syria	-ve ins.
Turkey	**+ve sig.**
Uganda	**+ve sig.**
Vietnam	-ve sig.

Figure 1
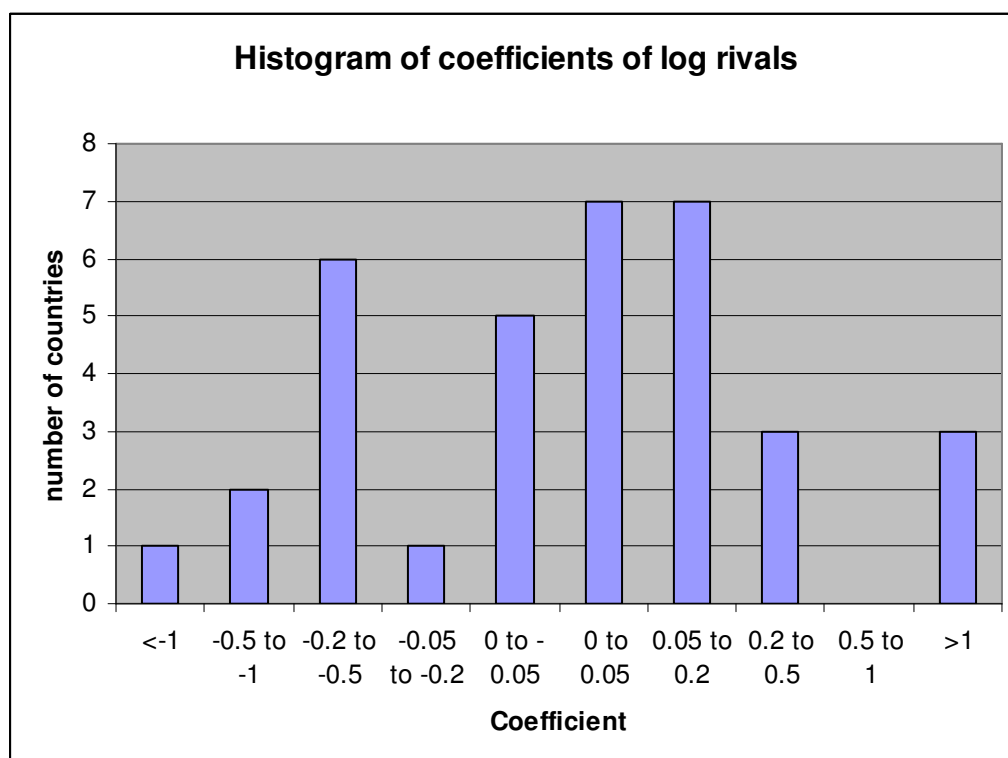
5. Conclusions

This paper has contributed the growing literature on the determinants of military spending, that uses cross-section and panel data to estimate models that incorporate 'arms-race' type effects. It has considered the theoretical and econometric questions that arise from Richardsonian action-reaction arms races between countries in terms of military expenditure, and that demonstrate military expenditure spillover effects across countries, in cross-section and panel data models. A key question is whether it is meaningful to talk of an 'arms race' in

panel data or cross-section data, and it has been suggested that such a concept is problematical. Rather, it may be more appropriate to talk about the relevant variables – aggregate military spending of the 'Security Web' (i.e. all neighbours and other security-influencing powers) and the aggregate military spending of 'Potential Enemies' or 'Rivals'. A number of econometric issues that arise in estimating and interpreting these models were also considered and illustrated by updating the model of the demand for military spending in Dunne and Perlo-Freeman (2003b) on SIPRI military expenditure data for the period 1988-2003. Broadly similar results were found, although the effect of neighbours' military spending is seen to come more equally from all the countries in a given country's Security Web, rather than particularly coming from rivals' military spending.

Comparing the results across different panel data estimation methods showed considerable variation. It was clear that the simple pooled OLS provides somewhat misleading results, while there were also differences between the fixed effects and random effects. The within and between estimators are likely to be measuring different things and so the differences in results were expected, but what was interesting that the cross section finds no role for economic variables and this is consistent with other cross section studies. Introducing lagged dependent variable to the fixed effects models showed the importance of recognising the dynamics of the processes.

While the Fixed Effects estimator does not allow for parameter heterogeneity in the slope coefficients, only in the intercepts, a random coefficients (mean group) estimator is precluded by the limited number of years of data. Although the FE estimator may nonetheless give consistent estimates of the average coefficients with I(1) variables (Dunne and Smith, 2007), the nature and importance of heterogeneity in the panel was investigated, by dividing the sample into 5 groups based on income. The results showed no obvious pattern to the differences and did not suggest that differences in income are the basis of the heterogeneity. A regional sub-classification sheds no more light, other than obvious points such that it was in the Middle East that the Iraq crises were most relevant, and in NATO and the Warsaw Pact during the Cold War. This means that while the panel results can suggest general tendencies that must apply in a majority of countries, the issues that are pertinent at a given time, and the way they impact, will vary considerably from country to country. When country-specific Rivals/Potential Enemies military spending variables are constructed, the expected presence of considerable heterogeneity was confirmed.

These findings suggest that the highly heterogenous pattern of determinants of military spending, means even considering two mutually hostile powers, their military burdens will depend on the whole nature of the relationship between them and other countries and events in the region, and not just a dyadic Richardsonian Action-reaction process. Thus the policy implications are that focussing too much on dyadic country interactions in attempting to limit arms races may fail to deal with the security problem and that wider regional security agreements may be necessary.

References

Paul Collier & Anke Hoeffler, 2007. "Unintended Consequences: Does Aid Promote Arms Races?," *Oxford Bulletin of Economics and Statistics*, Department of Economics, University of Oxford, vol. 69(1), pages 1-27, 02.

Collier, Paul and Hoeffler, Anke (2004) "Military Expenditure: Threats, Aid, and Arms Races". *Mimeo, University of Oxford*.

Dunne, J.P., Perlo Freeman, S. and Smith, R. (2008) "The Demand for Military Spending: Hostility versus Capability". *Defence and Peace Economics*, Vol. 19, no. 2, 2008, pp293-302.

Dunne, J.P. and Smith, R. (2007), "The Econometrics of Military Arms Races", Chapter 10 in (Keith Hartley and Todd Sandler eds.) *Handbook of Defence Economics*, North-Holland, 2007.

Dunne J. Paul, Eftychia Nikolaidou and Ron Smith (2003) Arms Race Models and Econometric Applications, in *Arms Trade, Security and Conflict* edited by Paul Levine and Ron Smith, Routledge p178-187.

Dunne, JP and Perlo Freeman, S (2003a) "The Demand for Military Spending in Developing Countries". *International Review of Applied Economics*, Vol. 17, no. 1, 2003, pp. 23-48.

Dunne, JP and Perlo Freeman, S (2003b) "The Demand for Military Spending in Developing Countries: A Dynamic Panel Analysis". *Defence and Peace Economics,* Vol 14, No. 6, 2003, pp. 461-474.

Fuertes A-M and R.P. Smith (2005) Panel Time Series, Cemmap course notes.

Murdoch, J.C. and Sandler, T (2002) Economic growth, civil wars, and spatial spillovers, *Journal of Conflict Resolution*, 46 91-110.

Goldsmith, BE (2003) "Bearing the Defence Burden, 1886-1989", *Journal of Conflict Resolution*, Vol . 47, No. 5, pp 551-573.

Murdoch, J.C. and Sandler, T (2004) Civil Wars and Economic Growth: Spatial Dispersion, American Journal of Political Science 48(1), 138-151.

Oren, I. (1994): "The Indo-Pakistani Arms Competition – a Deductive and Statistical Analysis", *Journal of Conflict Resolution*, Vol. 38 no. 2 pp 185-214

Polachek, S., Seiglie, C. and Xiang, J. (2005) "Globalization and International Conflict: Can FDI Increase Peace?", Rutgers University Newark Working Paper #2005-004.

Rosh, R. M. (1988): 'Third World Militarisation: Security Webs and the States they Ensnare', *Journal of Conflict Resolution* vol.32 no.4, pp 671-698.

Richardson, L. F. (1960): A*rms and Insecurity: A Mathematical Study of Causes and Origins of War.* Pittsburgh: Boxwood Press.

Smith, R. P . Sola, M and Spagnolo, F (2000) "The Prisoner's Dilemma and Regime-Switching in the Greek-Turkish Arms Race", *Journal of Peace Research*, 37, 6 pp 737-750.

Smith, R.P. and A. Tasiran (2005) The demand for arms imports, Journal of Peace Research, 42(2), 167-182.