



Does military expenditure crowd out health-care spending? Cross-country empirics

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Abstract

The trade-off between military expenditure and public health spending has remained an unsettled empirical issue. This paper investigates whether military expenditure has crowded out public health spending in 116 countries (including a subsample of 87 non-OECD countries) over the period 2000–2017. Through our system generalized methods of moments (GMM) estimations, we find that military expenditure, whether it is measured on a per-capita basis or as a proportion of total government expenditure, has a positive impact on the demand for health care. Nonetheless, we find a significant crowding-out effect of military expenditure on domestic government health spending by taking into account government fiscal capacity. The evidence we present supports the long-standing view that military expenditure has a particular ability to compete government financial resources away from publicly funded health spending. By interacting the military expenditure variable with income per capita, we find that an increase in income per capita has neutralized the crowding-out effect of military expenditure on domestic government health spending – less well-off countries stand to suffer most, and wealthy ones stand to suffer least, from the crowding-out effect. The crowding-out effect is statistically more specific to middle- and low-income countries in our samples.

Keywords Crowding-out effect · Opportunity costs · System GMM estimation · Trade-off

JEL classification C23 · H51 · O57

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1 Introduction

The potential trade-off between military expenditure and public health spending, which can be traced back to the opportunity costs of defence (Russett 1969; Dabelko and McCormick 1977), has been the recurrent subject of debate among analysts specializing in the fiscal burden for rising government expenditures. It is often assumed that military spending has opportunity costs and may crowd out other forms of public spending. The extent and form of crowding-out following an increase in military spending will depend on prior utilization and how the increase is financed. The fixed government fiscal capacity requires that an increase in military expenditure be financed by cuts in other public spending among other austerity measures. The trade-off between military expenditure and public health spending may occur when an increase in the former is crowding out an equivalent amount of the latter from total government expenditure. Tradeoffs are mostly measured in terms of budgetary expenditures. If the budget remains stable between years, budgetary decisions involve zero-sum games (Peroff and Podolak-Warren 1979). However, since total government expenditure typically increases from year to year, it is an increasing-sum rather than a zero-sum game (Harris et al. 1988). A negative relationship may not exist between the levels of military expenditure and other public spending, if the allocation process is an increasing-sum game. Negative shifts in the percentage allocation figures do not necessarily entail negative shifts in the levels (Peroff and Podolak-Warren 1979). Accordingly, the crowding-out effect of military expenditure on other public spending has been estimated based on ratios (i.e., both expenditures as proportions of total government expenditure or government fiscal capacity) rather than on absolute amounts (Peroff and Podolak-Warren 1979; Harris et al. 1988; Kollias and Paleologou 2011).

The past few years have seen a renewed interest in conducting cross-country empirical studies on the trade-off between military and public health expenditures. Lin et al. (2015) found a significant positive impact of military expenditure on health-care spending in a sample of 29 OECD countries over the period 1988–2005. Fan et al. (2018) found that increased military spending has a significant negative impact on both the publicly financed and total health expenditures of 197 countries over the period 2000–2013. Coutts et al. (2019) found that military burden has no significant impact on health burden in 18 countries in the Middle East and North Africa region over the period 1995–2011. Biscione and Caruso (2021) found that the once-lagged military expenditure–total government expenditure ratio has no significant effect on current health expenditure in a sample of 26 transition countries over the period 1990–2015. These studies differ widely in method and focus, and their empirical results point to rather different conclusions. Surprisingly, none of these recent studies has addressed the crowding-out effect of military expenditure on public health spending by taking into account government fiscal capacity (i.e., treating both expenditures as proportions of total government expenditure).

This caveat notwithstanding, it is still possible that the models of health-care expenditure tested in these studies were tainted with ad hocery in the choice of explanatory variables. The determinants of military- and health spending are quite distinct from each other (largely because military- and health spending decisions are made separately) and therefore should not enter the same demand equation. Empirically the military burden of a sovereign state is a function of the country's GNP, population, strategic status (i.e., interstate and civil conflict, whether the country is in the Middle East, etc.), regime type, and the aggregate

military spending of the country's neighbors and rivals (Dunne and Perlo-Freeman 2003). On the other hand, the macro demand for health care (i.e., per-capita health-care spending) is hypothesized to be a function of income per capita and non-income variables including the share of public expenditure in health care, population age structure, and the ratio of a health services price index to the GDP deflator (Hansen and King 1996).

The purpose of this paper is to investigate whether military expenditure has crowded out public health spending in 116 countries over the period 2000–2017. We address the question by asking how much an increase in military expenditure (i.e., per capita or as a proportion of total government expenditure) may affect per-capita health expenditure from a health economics perspective. We examine the tradeoffs through a macro health-care demand model assuming that military expenditure has a particular ability to compete government financial resources away from public health spending (Harris et al. 1988). Given government fiscal capacity limitation, the crowding-out effect is addressed by testing for the potential trade-off between military expenditure and domestic government health expenditure (both as proportions of total government expenditure). In addition, we test whether it is the interaction of military expenditure and income per capita that accounts for the cross-country variation in domestic government health spending. We further test for variations in these spending trade-offs over a subsample of 87 non-OECD countries, considering the heterogeneity of cross-country data.

The contribution of this paper is threefold. First, we provide new cross-country evidence on the crowding-out effect of military expenditure on public health spending. Second, this study represents a substantial revision to the methodological approach taken by previous studies in that we take into account government fiscal capacity when examining the crowding-out effect. Third, our results constitute a significant departure from previous studies by highlighting the fact that the crowding-out effect is statistically more specific to middle- and low-income countries.

The remainder of this paper is organized as follows. Section 2 reviews the literature. Section 3 specifies the empirical models and the data used. Section 4 explains the estimation methods and presents the results with a brief discussion. Section 5 summarizes and concludes.

2 Literature review

A neoclassical model of the state as a rational actor maximizing social welfare subject to government fiscal capacity is the starting point for an empirical analysis of the trade-off between military expenditure and public health spending. In the neoclassical theoretical framework, the state is in a position to determine a social welfare function through which military expenditure is determined fiscally by balancing its opportunity cost vis-à-vis education, public health, and other expenditures (Russett 1969; Dabelko and McCormick 1977; Peroff and Podolak-Warren 1979; Harris et al. 1988). The share of military expenditure is therefore an implicit function of national income, government fiscal capacity, and other socially important spending. One may well question about the rationality of actors assumed in formal neoclassical models. In some countries, military expenditure is independent of economic conditions and is predetermined by the internal logic of the state (Dunne and Perlo-Freeman 2003), irrespective of the budgetary discipline. A grand theory encompass-

ing the fiscal institutional realities of world countries is not available. On the other hand, we are not in a position to ignore the trade-offs between military expenditure and other socially important spending and public health spending in particular. Faced with a paucity of a universally adoptable theoretical framework, we have to settle for second best by taking a less formal modelling approach. In the existing literature, the Hansen and King (1996) health expenditure modeling approach has made an important contribution to empirical health-care expenditure studies. We introduce the military expenditure variable into the framework of Hansen and King (1996) to target on “the remaining unexplained cross-country variation in health expenditure” (Baltagi and Moscone 2010). In doing so, we tackle the potential trade-off between military expenditure and public health spending.

In recent literature, cross-country studies on the trade-off between military and public health expenditures have taken the lead in the field.¹ Lin et al. (2015) postulated a direct trade-off between military expenditure and public health spending. However, their results showed a significant, positive contemporaneous impact of military expenditure on health-care spending in a sample of 29 OECD countries over the period 1988–2005. Lin et al. (2015) argued that public health spending goes hand in hand with military expenditure because OECD countries are generally supportive of social welfare programs. Fan et al. (2018) assumed that health, military, and other government expenditures are jointly determined in the whole budget. Their results show that increased military burden (i.e., military expenditure as a percentage of GDP) has a significant, negative contemporaneous impact on both the publicly financed and total health expenditures (as a percentage of GDP) of 197 countries over the period 2000–2013. Fan et al. (2018) argued that increased military spending is a population health risk. Coutts et al. (2019) found that military burden has no significant impact on health burden in 18 countries in the Middle East and North Africa region over the period 1995–2011. Their empirical test for the “guns-vs.-butter” hypothesis was based on a tri-variate panel vector autoregressive model – the sample countries’ total health expenditure and total military expenditure, both as a percentage of GDP, and the sample countries’ casualties (i.e., deaths and serious injuries) due to terrorist attacks. Health burden was regressed on its own lagged value, military burden, and casualties. Their results showed that both military burden and casualties have a negative, insignificant contemporaneous impact on health burden. A null result may call into question the model specification of Coutts et al. (2019). Biscione and Caruso (2021) found that the once-lagged military expenditure–total government expenditure ratio has no significant effect on current health expenditure in a sample of 26 transition countries over the period 1990–2015. In spite of their inconclusive empirical evidence, Biscione and Caruso (2021) contributed to the “crowding-out effect” debate by their attempting investigating a delayed impact.

3 Methodology and data

3.1 The empirical models

There is no obvious theoretical prior when it comes to the impact of military expenditure on health-care spending. From a health economics perspective, income per capita is the most

¹ We will not go to great lengths to discuss those dated cross-country studies on this topic that involve the Cold-war factors. We refer interested readers to Fan et al. (2018) for a literature survey.

important factor explaining cross-country health expenditure differences. Additionally, the extent to which health expenditure is financed by the government and population age structure have been flagged as two important factors in explaining cross-country variation in health expenditure. The remaining cross-country variation in health expenditure has been left largely unexplained (Baltagi and Moscone 2010). As Peroff and Podolak-Warren (1979) said in their seminal article, there are opportunity costs to military spending because anticipated increases in military spending constrain policymakers from increasing non-military spending and introducing new programs in health. Considering government fiscal capacity limitation, we argue that military expenditure would affect health-care spending and formulate a dynamic panel model as follows:

$$\log(h_{it}) = \beta_1^1 \log(h_{i,t-1}) + \beta_2^1 \log(m_{i,t-j}) + \beta_3^1 \log(pub_{it}) + \beta_4^1 \log(y_{it}) + \beta_5^1 \log(old_{it}) + \beta_6^1 \log(yng_{it}) + \mu_i + \lambda_t + \epsilon_{it} \quad (1)$$

$$\log(h_{it}) = \beta_1^2 \log(h_{i,t-1}) + \beta_2^2 \log(rm_{i,t-j}) + \beta_3^2 \log(pub_{it}) + \beta_4^2 \log(y_{it}) + \beta_5^2 \log(old_{it}) + \beta_6^2 \log(yng_{it}) + \mu_i + \lambda_t + \epsilon_{it} \quad (2)$$

where the subscript i denotes the i th country ($i = 1, \dots, N$) and t denotes the t th year ($t = 1, \dots, T$). Here h stands for real per-capita health expenditure. By including the lagged dependent variable in the equations, we model the dynamics of the spending process to allow for a hangover from the previous expenditure on health. Here m and rm stand for real per-capita military expenditure and the military expenditure–total government expenditure ratio (hereafter as “military expenditure ratio”), respectively; by setting $j = 0, 1$ we consider the military expenditures in year t or alternatively in year $t - 1$.² Here pub stands for public health coverage (i.e., the proportion of health expenditure that is publicly funded; current military expenditure may affect subsequent pub and therefore has health-care spending consequences), y is real GDP per capita, and old and yng denote the populations aged 65 and over and 0–14 as percentages of working-age population, respectively. These last two variables represent the different health-care needs of different age groups. The disturbance term is specified as a two-way error component model where μ denotes a country-specific effect, λ denotes a year-specific effect, and ϵ is the remainder disturbance. The country-specific effects represent those characteristics that are peculiar to the sample countries, including the presence of a centralized national health system. The time-period effects are assumed to be fixed parameters to be estimated as coefficients of time dummies for each period in the sample. We use time dummies as a proxy for the underlying causes of the rise and fall in health-care spending, e.g., medical care technology progress, policy shifts, and new diseases. Without adding a time-specific effect to the models in question, potential structural breaks were left unaccounted for, giving rise to biased estimates.³

Assuming that military expenditure has a particular ability to compete government financial resources away from publicly funded health spending due to government fiscal capacity

²By lagging the military expenditure variable once, we address the probable simultaneity or reverse causality between the dependent variable and the military expenditure variable.

³Moreover, the Arellano–Bond autocorrelation test and the robust estimates of the coefficient standard errors in GMM estimations assume no contemporaneous correlation in the idiosyncratic disturbances. Time dummies make this assumption more likely to hold (Roodman 2009).

limitation, we examine the crowding-out effect by testing for the potential trade-off between military expenditure and domestic government health spending (i.e., both expenditures as proportions of total government expenditure) as follows:

$$\log(rg_{it}) = \beta_1^3 \log(rg_{i,t-1}) + \beta_2^3 \log(rm_{i,t-j}) + \beta_3^3 \log(pub_{it}) + \beta_4^3 \log(y_{it}) + \beta_5^3 \log(old_{it}) + \beta_6^3 \log(yn_{it}) + \mu_i + \lambda_t + \epsilon_{it} \quad (3)$$

where $j = 0, 1$, and rg stands for the domestic government health expenditure–total government expenditure ratio (hereafter as “domestic government health expenditure ratio”).

In public expenditure terms, wealthy governments may spend more than their less well-off counterparts do on both health care and defence – hence the patterns of the potential trade-off between military expenditure and domestic government health spending may vary across countries. In view of this possibility, we carry the estimations a stage further by interacting the military expenditure ratio with real GDP per capita as follows:

$$\log(rg_{it}) = \beta_1^4 \log(rg_{i,t-1}) + \beta_2^4 \log(rm_{i,t-j}) + \beta_3^4 \log(pub_{it}) + \beta_4^4 \log(y_{i,t-j}) + \beta_5^4 \log(old_{it}) + \beta_6^4 \log(yn_{it}) + \beta_7^4 \left\{ \log(rm_{i,t-j}) \times \log(y_{i,t-j}) \right\} + \mu_i + \lambda_t + \epsilon_{it} \quad (4)$$

where $j = 0, 1$. The estimated coefficient on the interaction term would indicate whether it is the combination of the military expenditure ratio and real GDP per capita that “explains” the cross-country variation in the domestic government health expenditure ratio.

3.2 Estimation methods

In the models we specified, the dynamic relationships are characterized by the presence of a lagged dependent variable among the regressors. The lagged dependent variable is correlated with the error term by construction, and this renders the OLS estimator biased and inconsistent even if the remainder disturbances are not serially correlated. The fixed-effects within estimator is biased of $O(1/T)$ (Nickell 1981) and is inconsistent for “large N , small T ” panels (Baltagi 2013, 155). Anderson and Hsiao (1982) suggested first-differencing the model to get rid of the individual-specific effect and using the (first-differenced) twice-lagged dependent variable as an instrument for the first-differenced lagged dependent variable. The Anderson-Hsiao estimator leads to consistent but not necessarily efficient estimates of the parameters in the model, because it does not make use of all the available moment conditions and does not take into account the differenced structure on the residual disturbances (Baltagi 2013, 156). Arellano and Bond (1991) proposed a GMM procedure that is more efficient than the Anderson-Hsiao estimator by obtaining additional instruments through the orthogonality conditions that exist between lagged values of the dependent variable and the differenced errors (Baltagi 2013, 156, 157). A special feature of dynamic panel GMM estimation is that the number of moment conditions increases with T . In system GMM estimation, the quadratic growth of moment conditions with respect to T causes “instrument proliferation”. To keep the instrument count below N , we use only certain lags of all available lags for instruments or “collapse” them into smaller sets (Roodman 2009). To ensure the joint validity of the instruments, a post-estimation Sargan/Hansen test for the over-identification restrictions is needed. We perform the Hansen test because the Sargan

test requires that the errors are homoskedastic for the sake of consistency. To ensure the consistency of their GMM estimator, Arellano and Bond (1991) proposed a test for the absence of second-order serial correlation for the disturbances of the first-differenced equation (Baltagi 2013, 160). In all our estimations reported below, the Hansen test for over-identification does not reject the null (indicating that the instruments are jointly valid), the AR(1) test indicates first-order serial correlation in the disturbances, and the AR(2) test indicates that the differenced errors are serially uncorrelated. As such, the basic identification assumptions of the system-GMM equations are valid.

Blundell and Bond (1998) suggested a system GMM estimator using extra moment conditions with first-differenced instruments for the equation in levels and instrument in levels for the first-differenced equation, because the first-differenced GMM estimator applied to short- T and highly persistent panel data may suffer from a severe small-sample bias due to weak instruments. The system GMM estimator is designed for “large N , small T ” panels that contain fixed effects and idiosyncratic disturbances that are heteroskedastic and correlated within (but not across) individuals (Roodman 2009). Through their simulations, Blundell et al. (2001) found that the system GMM estimator not only improves the precision but also reduces the finite-sample bias when the standard first-differenced estimator gives rise to large finite-sample bias and very low precision for the parameters in dynamic panel models with weakly exogenous covariates. In contrast to the first differenced-GMM estimator, the system GMM estimator based on the level-equation moment condition (Blundell and Bond 1998) is consistent and asymptotically normal even under the unit root hypothesis (Westerlund and Breitung 2013). In our models, apart from the lagged dependent variable, the other regressors might be weakly endogenous with respect to the idiosyncratic error component. We proceed to rely on the system GMM estimator to address the problem of weakly endogenous regressors. We derive the system GMM estimates by using the Stata command *xtabond2* (Roodman 2009) which performs the Windmeijer (2005) finite-sample correction to the standard errors in two-step GMM estimation and small-sample adjustments.

3.3 Data description

We use a balanced panel dataset of 116 countries (Table A1) over the period 2000–2017 ($T=18$, subject to data availability), including 29 countries that formed the Organization for Economic Co-operation and Development (OECD) in the year 2000. This sample excludes those countries where a substantial portion of the data are missing due to intra-state (inter-state) conflicts, regime collapses, or low transparency. The data on health expenditure and public health coverage were obtained from the WHO Global Health Expenditure Database that has recorded the member countries’ health expenditure data from 2000 and on. The data on military expenditure were obtained from the SIPRI Extended Military Expenditure Database. The current-price expenditure figures we obtained from these two databases were originally derived by using market exchange rates. We followed the practice of the World Bank World Development Indicators (WDI) and converted those current-price U.S. dollar figures into constant 2010 U.S. dollars. The data on GDP per capita and the ratios of older and younger dependents were sourced from the WDI. Table A2 shows the definition of each

of the variables. Table A3 provides the descriptive statistics. All observations that we use to estimate the equations in Sect. 3.1 are three-year averages.⁴

4 Results and discussion

4.1 Per-capita health expenditure

From Table 1, the estimated coefficient on the lagged dependent variable is positive and significant, showing that per-capita health expenditure of the full sample has stronger persistence (0.728–0.661) than that of the non-OECD sample (0.557–0.563). From columns

Table 1 Per-capita health expenditure, I

	Full sample		Non-OECD sample	
	(1)	(2)	(3)	(4)
Lagged dep. variable	0.728*** (0.073)	0.661*** (0.081)	0.557*** (0.102)	0.563*** (0.129)
Mil. ex. per capita	0.08* (0.043)		0.095** (0.047)	
Mil. ex. per capita $_{t-1}$		0.038 (0.045)		0.136** (0.066)
Pub. health coverage	0.048 (0.06)	0.037 (0.13)	0.037 (0.057)	0.023 (0.122)
Real GDP per capita	0.04 (0.087)	0.198* (0.107)	0.132 (0.107)	0.069 (0.147)
Age dep. ratio, old	0.116** (0.053)	0.053 (0.087)	0.03 (0.066)	0.125 (0.078)
Age dep. ratio, young	-0.283*** (0.091)	-0.25** (0.11)	-0.427*** (0.104)	-0.329* (0.174)
Observations	580	580	435	435
Countries	116	116	87	87
Instruments	78	73	78	73
Hansen test	0.265	0.411	0.409	0.389
AR(1) test	0.001	0.004	0.011	0.027
AR(2) test	0.667	0.574	0.397	0.776

The dependent variable is the log of per-capita health expenditure. All variables are in natural logarithm. The estimated coefficients and standard errors are obtained from the system GMM two-step estimation with the Windmeijer correction and small-sample adjustments. A set of year effects and a constant are included in all specifications. Robust standard errors in parentheses. The p -values for the Hansen test and the Arellano-Bond AR tests are reported.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

⁴To minimize the time dimension of the data, we transformed the data into three-year averages (with six individual time periods) because the Arellano and Bond (1991) test for both AR(1) and AR(2) based on the residuals from the two-step estimator of the first-differenced equation can be performed only on samples where $T \geq 5$.

(1) and (3) the estimated coefficients on per-capita military expenditure are 0.08 and 0.095 in elasticity, respectively. From column (4) the estimated coefficient on the lagged per-capita military expenditure is 0.136 in elasticity. Accordingly, the estimated cumulative effects (which is given by $\beta_2^1/(1-\beta_1^1)$ in Eq. (1)) are 0.294 in column (1) and 0.214–0.311 in columns (3) and (4), respectively. These estimated coefficients indicate a positive impact of military expenditure on health-care spending in both the full sample and non-OECD sample. From column (2) the estimated coefficient on real GDP per capita is 0.198 in elasticity. From column (1) the estimated coefficient on the ratio of older dependents is 0.116 in elasticity. The estimated coefficients on the ratio of younger dependents range from -0.283 in column (1) to -0.25 in column (2) in elasticity; they range from -0.427 in column (3) to -0.329 in column (4) in elasticity. A positive (negative) coefficient on the ratio of older (younger) dependents suggests that the elderly (young) consume more (less) health care per capita than the working age (Hansen and King 1996).

From Table 2, the estimated coefficient on the lagged dependent variable is positive and significant, showing that per-capita health expenditure of the full sample has stronger persistence (0.732–0.738) than that of the non-OECD sample (0.634–0.63). From column (3) the estimated coefficient on the military expenditure ratio is 0.108 in elasticity. Accordingly, the estimated cumulative effect is 0.295 in column (3). The estimated coefficients on real GDP per capita range from 0.134 in column (1) to 0.129 in column (2) in elasticity; they range

Table 2 Per-capita health expenditure, II

	Full sample		Non-OECD sample	
	(1)	(2)	(3)	(4)
Lagged dep. variable	0.732*** (0.055)	0.738*** (0.057)	0.634*** (0.072)	0.63*** (0.087)
Mil. ex. ratio	0.042 (0.049)		0.108* (0.065)	
Mil. ex. ratio $_{t-1}$		0.011 (0.05)		0.071 (0.068)
Pub. health coverage	0.052 (0.121)	0.052 (0.121)	0.06 (0.133)	0.049 (0.116)
Real GDP per capita	0.134** (0.068)	0.129* (0.072)	0.169* (0.094)	0.167* (0.096)
Age dep. ratio, old	0.149*** (0.046)	0.126** (0.05)	0.111** (0.052)	0.088 (0.077)
Age dep. ratio, young	-0.168 (0.104)	-0.184* (0.096)	-0.27** (0.135)	-0.302** (0.141)
Observations	580	580	435	435
Countries	116	116	87	87
Instruments	78	73	78	73
Hansen test	0.332	0.506	0.332	0.326
AR(1) test	0.001	0.001	0.003	0.005
AR(2) test	0.655	0.686	0.501	0.609

See Table 1.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

from 0.169 in column (3) to 0.167 in column (4) in elasticity. These estimated coefficients indicate a positive, significant impact of income per capita on health-care spending in both the full sample and non-OECD sample.

4.2 The crowding-out effect

From Table 3, the estimated coefficient on the lagged dependent variable is positive and significant, showing that the domestic government health expenditure ratio of the full sample has stronger persistence (0.749–0.783) than that of the non-OECD sample (0.716–0.746). From columns (1) and (3) the estimated coefficients on the military expenditure ratio are -0.104 and -0.163 in elasticity, respectively. From columns (2) and (4) the estimated coefficients on the lagged military expenditure ratio are -0.127 and -0.168 in elasticity, respectively. Accordingly, the estimated cumulative effects range from -0.414 in column (1) to -0.585 in column (2), and they range from -0.574 in column (3) to -0.661 in column (4). These estimated coefficients indicate a negative, significant impact of the military expenditure ratio on the domestic government health expenditure ratio in both the full sample and non-OECD sample. The estimated coefficients on public health coverage range from 0.119 in column (1) to 0.118 in column (2) in elasticity; from columns (3) and (4) both coefficients are 0.158 in elasticity.

Table 3 Crowding-out effect

	Full sample		Non-OECD sample	
	(1)	(2)	(3)	(4)
Lagged dep. variable	0.749*** (0.084)	0.783*** (0.086)	0.716*** (0.088)	0.746*** (0.085)
Mil. ex. ratio	-0.104^* (0.059)		-0.163^{**} (0.074)	
Mil. ex. ratio $_{t-1}$		-0.127^{**} (0.058)		-0.168^{**} (0.068)
Pub. health coverage	0.119* (0.063)	0.118** (0.058)	0.158** (0.081)	0.158** (0.074)
Real GDP per capita	0.007 (0.032)	0.006 (0.026)	0.013 (0.043)	0.006 (0.038)
Age dep. ratio, old	0.029 (0.053)	0.062 (0.059)	0.023 (0.067)	0.059 (0.073)
Age dep. ratio, young	-0.069 (0.085)	-0.073 (0.091)	-0.008 (0.13)	-0.027 (0.133)
Observations	580	580	435	435
Countries	116	116	87	87
Instruments	78	73	78	73
Hansen test	0.429	0.408	0.644	0.532
AR(1) test	0.042	0.042	0.037	0.035
AR(2) test	0.163	0.247	0.193	0.32

The dependent variable is the log of domestic government health expenditure ratio.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

In order to provide evidence on heterogeneity in our findings, we interact the two terms in Eq. (4), which allows us to examine whether the crowding-out effect varies across countries of different income levels. We expect an increase in income per capita to have a positive impact on the domestic government health expenditure ratio, so a positive (negative) estimated coefficient on the interaction term would indicate that the military expenditure ratio in interaction with this indicator of income level is associated with a weaker (stronger) crowding-out effect. Where the interaction term is significant, the estimated coefficients on $\log(rm_{i,t-j})$ and $\log(y_{i,t-j})$ cannot be interpreted in the conventional way. Instead, the partial derivative $\beta_2^4 + \beta_7^4 \log(y_{i,t-j})$ is evaluated at the mean, minimum and maximum values of real GDP per capita, because this derivative varies within the sample depending on the magnitude of real income per capita.

From Table 4 columns (1) to (4), the estimated coefficients on the military expenditure ratio (current and lagged) are significant. Importantly, the estimated coefficients on the inter-

Table 4 Interaction effects

	Full sample		Non-OECD sample	
	(1)	(2)	(3)	(4)
Lagged dep. variable	0.779*** (0.086)	0.831*** (0.078)	0.768*** (0.081)	0.828*** (0.075)
Mil. ex. ratio	-1.286*** (0.492)		-1.145** (0.515)	
Mil. ex. ratio $_{t-1}$		-1.076*** (0.337)		-1.15*** (0.399)
Pub. health coverage	0.107** (0.054)	0.057 (0.064)	0.141** (0.066)	0.059 (0.08)
Real GDP per capita	0.207*** (0.079)		0.237** (0.108)	
Real GDP per capita $_{t-1}$		0.154* (0.079)		0.191* (0.106)
Age dep. ratio, old	0.147* (0.081)	0.123** (0.059)	0.138* (0.078)	0.134* (0.069)
Age dep. ratio, young	-0.143 (0.108)	-0.142 (0.096)	-0.094 (0.102)	-0.142 (0.097)
{Mil. ex. ratio \times Real GDP per capita}	0.128*** (0.049)		0.117** (0.055)	
{Mil. ex. ratio $_{t-1} \times$ Real GDP per capita $_{t-1}$ }		0.108*** (0.035)		0.118*** (0.044)
Observations	580	580	435	435
Countries	116	116	87	87
Instruments	91	86	83	78
Hansen test	0.314	0.335	0.328	0.359
AR(1) test	0.035	0.035	0.039	0.04
AR(2) test	0.164	0.267	0.154	0.291

See Table 3.

* Significant at the 10% level.

** Significant at the 5% level.

*** Significant at the 1% level.

action term are positive and significant. There is now evidence that real GDP per capita in interaction with the military expenditure ratio is associated with higher levels of domestic government health spending. The negative relationship between the military expenditure ratio and the domestic government health expenditure ratio diminishes – becomes more positive, in fact – the higher real GDP per capita there is. Using the estimated coefficients in column (1), the partial derivative of the domestic government health expenditure ratio with respect to the military expenditure ratio at the mean level of real GDP per capita is -0.09 . The same derivative evaluated at the minimum level of real GDP per capita is -0.46 . When real GDP per capita is at its maximum value, the derivative takes a positive value of 0.18 . Column (3) reports the results when all OECD countries are excluded. Using the estimated coefficients in column (3), the partial derivative at the mean level of real GDP per capita is -0.094 . The same derivative evaluated at the minimum level of real GDP per capita is -0.388 . When real GDP per capita is at its maximum value, the derivative takes a positive value of 0.197 . Using the estimated coefficients in columns (2) and (4) of Table 4, we evaluate the partial derivative of the domestic government health expenditure ratio with respect to the lagged military expenditure ratio, then again our results hardly change in qualitative terms. The above results suggest that an increase in real GDP per capita has neutralized the crowding-out effect of military expenditure on domestic government health spending – less well-off countries stand to suffer most, and wealthy ones stand to suffer least, from the crowding-out effect.

Depending on the sign of the partial derivative $\beta_2^4 + \beta_7^4 \log(y_{i,t-j})$ in Eq. (4), the impact of military expenditure on domestic government health spending may be either positive or negative. By setting this partial derivative to zero, we solve for the threshold level of real income per capita using the equation $y_{i,t-j} = \exp(-\beta_2^4/\beta_7^4)$. Military expenditure has a statistically discernable crowding-out effect on domestic government health spending in countries with a real income per capita below this threshold level. Using the estimated coefficients in Table 4 columns (1) and (2), the threshold income levels are around 23 and 21 thousand constant 2010 U.S. dollars, respectively. Using the estimated coefficients in Table 4 columns (3) and (4), the threshold income levels are around 18 and 17 thousand constant 2010 U.S. dollars, respectively. The above results suggest that the crowding-out effect is statistically more specific to middle- and low-income countries. We discuss this finding in the section below.

4.3 Discussion

We find a positive, significant impact of per-capita military expenditure on per-capita health expenditure in the full sample, and we find a positive, significant impact of both the current and lagged per-capita military expenditures on per-capita health expenditure in the non-OECD sample. Additionally, we find a positive, significant impact of the military expenditure ratio on per-capita health expenditure in the non-OECD sample. The evidence we present above gives more weight to the hypothesis of positive effects of military expenditure on health-care spending rather than some zero-sum game type of tradeoffs. Nonetheless, we find a significant crowding-out effect of military expenditure on domestic government health spending in both the full sample and non-OECD sample by taking into account government fiscal capacity (i.e., both expenditures as proportions of total government expenditure). The evidence we present immediately above supports the long-standing view that

military expenditure has a particular ability to compete government financial resources away from publicly funded health spending.

Importantly, we find some, if limited, evidence that an increase in real income per capita has neutralized the crowding-out effect of military expenditure on domestic government health spending. As a result, less well-off countries stand to suffer most, and wealthy ones stand to suffer least, from the crowding-out effect. This finding is consistent with Fan et al. (2018), whose authors reached some similar conclusion by splitting their full sample. The threshold income levels we derived (Sect. 4.2) confirm that middle- and low-income countries are less able to minimize the crowding-out effect compared to their high-income counterparts. This finding is instrumental in resolving the dramatic opposition between a positive impact of military expenditure on health-care spending in 29 OECD countries (Lin et al. 2015) and a negative one in 197 countries (Fan et al. 2018) with highly varied income levels. The crowding-out effect becomes statistically more discernable as the sample consists mainly of countries with a real income per capita below the threshold level.

Throughout our estimations, the estimated coefficients on public health coverage, real GDP per capita and elderly (young) dependency ratio have *a priori* consistent signs, although the significance levels of the individual coefficients vary across specifications and samples.⁵

5 Conclusions

We look into whether military expenditure has crowded out public health spending in 116 countries over the period 2000–2017. First, we find that military expenditure, whether it is measured on a per-capita basis or as a proportion of total government expenditure, has a positive impact on the demand for health care. Second, we find a significant crowding-out effect of military expenditure on domestic government health spending by taking into account government fiscal capacity. Third, we find that less well-off countries stand to suffer most, and wealthy ones stand to suffer least, from the crowding-out effect. In this study, we underscore the importance of government fiscal capacity in disentangling the crowding-out effect of military expenditure on domestic government health spending from all the postulated trade-offs. Based on our evidence, we suggest that freeing up government financial resources that would be drained by the military for health-care spending is particularly relevant to the prospects for human development in low- and middle-income countries. Faced with the unprecedented challenge of surviving COVID-19, many countries have seen a sudden, unexpected rise in public health spending. In spite of the devastating pandemic, the biggest military spenders (including the largest developing economies) have again raised their annual defence budgets. Given that military expenditure may crowd out public health-care spending out of total government expenditure, this aberrant policy development is especially worrisome and merits further attention.

⁵We also performed estimations with government spending (as a percentage of GDP) and trade openness as additional control variables. However, the estimated coefficients on these two variables were insignificant. In public expenditure terms, democratic regimes spend more than their autocratic counterparts do on health care. However, it should be emphasized that the measure of regime type (e.g., the Polity2 score) lacks variation over time and the level effects of democracy may be absorbed by the country fixed effect.

One obvious extension of this paper is to allow for time-varying slope coefficients in the dynamic panel models of demand for health care to see if the magnitude of the trade-offs will become increasingly larger when the defence-spending share of total government expenditure grows. It would be worth investigating the patterns of such time-varying trade-offs across countries of different regions and income groups. We hope to pursue that line of research in the future when some informative “large N , large T ” cross-country panels become available.⁶

Appendix

Table A1 Sample countries ($N=116$)

OECD	Non-OECD		
Australia	Algeria	Fiji	Namibia
Austria	Angola	Georgia	Nepal
Belgium	Argentina	Ghana	Nicaragua
Canada	Azerbaijan	Guatemala	Nigeria
Czech Republic	Bahrain	Guyana	North Macedonia
Denmark	Bangladesh	Honduras	Oman
Finland	Belize	India	Pakistan
France	Bolivia	Indonesia	Papua New Guinea
Germany	Botswana	Iran, Islamic Rep.	Paraguay
Greece	Brazil	Israel	Peru
Hungary	Brunei Darussalam	Jamaica	Philippines
Ireland	Bulgaria	Jordan	Romania
Italy	Burkina Faso	Kazakhstan	Russian Federation
Japan	Cabo Verde	Kenya	Rwanda
Korea, Rep.	Cambodia	Kuwait	Saudi Arabia
Luxembourg	Cameroon	Kyrgyz Republic	Senegal
Mexico	Chad	Latvia	Serbia
Netherlands	Chile	Lebanon	Seychelles
New Zealand	China	Lesotho	Sierra Leone
Norway	Colombia	Lithuania	Singapore
Poland	Croatia	Madagascar	Slovenia
Portugal	Cyprus	Malaysia	Sri Lanka
Slovak Republic	Dominican Republic	Mali	Tanzania
Spain	Ecuador	Malta	Thailand
Sweden	Egypt, Arab Rep.	Mauritius	Trinidad and Tobago
Switzerland	El Salvador	Moldova	Tunisia
Turkey	Estonia	Mongolia	Uganda
United Kingdom	Eswatini	Morocco	Ukraine
United States	Ethiopia	Mozambique	Uruguay

⁶One reviewer called our attention to “asymmetric panel causality” (Hatemi-J et al. 2018) and “hidden panel cointegration” (Hatemi-J 2020). Both of the papers focused on a two-variable case. The asymmetric causality tests (Hatemi-J et al. 2018) dealt with individual time series rather than the whole panel. Hatemi-J (2020) concerned a “large T ” panel ($T=92$) comprised of three asymmetric time series. The time dimension of the present panel is too short ($T=18$) for investigating “asymmetric panel causality” or “hidden panel cointegration”. We reserve his suggestion for future research.

Table A2 Variables and data sources

Variable	Definition	Source
Health ex. per capita	Current health expenditure (current US\$) divided by total population	WHO
Domestic gov't health ex. ratio	Domestic government health expenditure as a percentage of general government expenditure	WHO
Mil. ex. per capita	All current and capital expenditures on the armed forces divided by total population (current US\$)	SIPRI
Mil. ex. ratio	Military expenditure as a percentage of general government expenditure	SIPRI
Public health coverage	Domestic government health expenditure and external health expenditure as a percentage of current health expenditure	WHO
Real GDP per capita	GDP (constant 2010 US\$) divided by total population	WDI
Age dep. ratio, old	The population aged 65 and over as a percentage of working-age population	WDI
Age dep. ratio, young	The population aged 0–14 as a percentage of working-age population	WDI

WHO Global Health Expenditure Database; SIPRI Extended Military Expenditure Database; World Bank World Development Indicators. We converted the current US\$ figures into constant 2010 U.S. dollars when preparing the dataset.

Table A3 Descriptive statistics (N=116)

Variable	Symbol	No. of obs.	Mean	SD	SD-Between	SD-Within	Min.	Max.
Health ex. per capita	<i>h</i>	2088	1357	1536.6	1513	300.6	24.9	9298
Domestic gov't health ex. ratio	<i>rg</i>	2088	10.5	4.2	3.85	1.64	1.3	23.6
Mil. ex. per capita	<i>m</i>	2088	434	734	719	160.6	6.8	6725
Mil. ex. ratio	<i>rm</i>	2088	6.7	5.3	4.9	2	0.6	33
Public health coverage	<i>pub</i>	2088	59.3	17.8	17.2	5	12.6	95
Real GDP per capita	<i>y</i>	2088	19528	18759.4	18636.7	2723.6	613	97864
Age dep. ratio, old	<i>old</i>	2088	13.6	8.4	8.3	1.5	2.3	45.1
Age dep. ratio, young	<i>yng</i>	2088	44.7	22	21.7	4.2	15.8	103.6

SD – standard deviation.

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Data & code availability The data are publicly accessible through the WHO Global Health Expenditure Database, the SIPRI Extended Military Expenditure Database, and the World Bank World Development Indicators. The code is available upon reasonable request.

Declarations

Conflict of interest The authors have no conflicts of interest to declare.

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