MODELING THE US-SOVIET ARMS RACE AS AN ERROR-CORRECTION PROCESS

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Abstract

If two states are engaged in an arms race, then we expect to find their armament levels moving together over time in equilibrium. Such a "moving equilibrium" in US and Soviet armament levels -- or *cointegration*, in econometric parlance -- represents a commonsense understanding of the superpower arms race. Most empirical investigations have started with this commonsense understanding and have sought to reveal the short-run action-reaction dynamic that generates this moving equilibrium in military expenditures. But to hypothesize such an action-reaction process is to expect rather much from less-than omniscient policymakers and their lumbering state bureaucracies, and the empirical literature offers little evidence to the contrary. The preoccupation with short-run dynamics also neglects the important long-run dimension of the arms race. If the minimum condition for the existence of an arms race is the co-movement of two states' military expenditures over time, then it must be the case that this long-run equilibrium is maintained by one or both. Time-series analysts have termed such behavior *error correction*.

Our analysis indicates that six pairs of time series -- Soviet military outlays and US service requests, presidential requests, and congressional authorizations in both current and constant dollars -- are cointegrated, suggesting that US and Soviet armament efforts maintained a long-run equilibrium relationship. That is the pattern we usually associate with an arms race. But the relationship was not symmetrical, for it was the US policymakers and not the Soviets who adjusted their defense budgets in the face of deviations from equilibrium. The superpower arms race appears from this evidence to have been one sided: the Soviets acted, the Americans reacted. Relative to these long-run dynamics, evidence for US reactivity in the short-run is sparse. Overall, the impact of Soviet outlays on US defense budgeting was mostly indirect -- i.e., to the extent that Soviet policy caused deviations in the equilibrium relationship, which US policymakers sought to correct.

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More than any other dimension of international relations, action-reaction processes (or "reciprocity") have occupied the modeling efforts of scholars committed to the scientific study of world politics. And of the action-reaction processes examined, surely no other has received more attention than the US-Soviet arms race. The preoccupation with the superpower arms race is partly due to the fact that it constituted a basic ingredient of the cold war. But the preoccupation also reflect scholarly frustration. We all know, or think we know, that during the cold war the superpowers were engaged in an arms race; yet it has been exceedingly difficult to demonstrate this via statistical analysis of military expenditure data. As Etcheson (1989: 3) has remarked: "The finding is that there is apparently no military interaction between the world's two dominant military powers. Such a conclusion is genuinely puzzling."

In this paper, we discuss the different empirical characteristics of an arms race, and show that there is in fact evidence of a US-Soviet arms race. This evidence is derived from a fairly straightforward statistical analysis of military expenditure data, and without recourse to the reconceptualizations, the use of control variables, or the sometimes tortured logic that scholars have found necessary in order to arrive at the same conclusion. In particular, we employ an error-correction approach, which is designed to capture action-reaction dynamics operating in the long run as well as the short run. The former have often been missed in the empirical arms-race literature to date. Our analysis sheds light not only on the short-run action-reaction processes typically investigated by arms-race researchers, but also the long-run equilibrium relationship between US and Soviet military expenditures and the process of adjustment in the face of disequilibrium.

We will not attempt a thorough review of the arms-race literature here since this has been so effectively done by others (see especially Etcheson 1989; Intriligator and Brito 1989; Isard and Anderton 1988). Instead, we begin with a brief overview of the theoretical notion of an arms race, as pioneered by Richardson, and the primary obstacles to operationalizing the arms-race model in empirical research. The Richardson framework is consistent with a reasonably sophisticated and multi-faceted conception of an arms race, richer than its either its critics or existing empirical research would seem to suggest. The error-correction methodology is truer to this framework than are other approaches in that the theoretical concepts of equilibrium and reequilibration correspond to the operational concepts of *cointegration* and *error correction* in time-series analysis. We detail a step-by-step procedure for testing the cointegration of time series and for estimating error-correction models, and then apply these tools in our investigation of action-reaction dynamics in US and Soviet military spending.

Unfortunately, despite the appropriateness of the time-series methodology employed here, we must equivocate in drawing conclusions about the nature of the US-Soviet arms race. The basic obstacle preventing a balanced examination of the process is the unavailability of Soviet defense budgeting data. It is during the defense budgeting process, and not at the point of actual expenditure, where we expect to find the strongest evidence of reactivity to the defense expenditures of a rival. Until data on the Soviet defense budgeting process become available for the cold war period, we must resign ourselves to essentially one-sided examinations of the superpower arms race.

THE ARMS RACE MODEL IN THEORY

Richardson's basic action-reaction model is by now familiar to arms-race modelers and non-modelers alike. He imagined a process in which one side in a competitive relationship reacts to increases in the other side's arms by increasing its own arms. The only restraint on this reaction is each side's realization that there are diminishing marginal returns on the investment in defense. For two rivals with armament levels *x* and *y*, Richardson (1960) gave the process this mathematical form:

$$\begin{aligned} dx/dt &= ky - ax + g \\ dy/dt &= lx - by + h \end{aligned} \qquad [1]$$

That is, the rate of change in one side's armament levels rises with increases in the other side's armaments (at a rate indicated by "defense coefficients" k and l), while it falls with increases in its own armaments (at the rate of a and b, the "fatigue coefficients"). The fact that these two states are in a competitive relationship to begin with is represented by the positive constants, or "grievance terms," g and h.

An equilibrium exists where neither side feels it necessary to continue arming -- i.e., where dx/dt = dy/dt = 0. Here, the reaction functions are:

$$x = k'y + g' \quad \text{where } k' = k/a \text{ and } g' = g/a \qquad [3]$$

$$y = l'x + h' \quad \text{where } l' = l/b \text{ and } h' = h/b \qquad [4]$$

The armament levels for each country, therefore, is a linear function of the arms held by the other (with defense coefficients k' and l' now deflated by fatigue coefficients a and b, respectively).[1] Assuming that the grievance terms g and h are positive, a stable equilibrium exists if $k' l' < 0$.[2] This is represented graphically in Figure 1a. The single line represents the minimum security requirements of one rival state, here the United States. The US wants to maintain its armaments at a level equal to or exceeding g' plus some portion k' of the arms held by its rival, the Soviet Union. The Soviet Union, as indicated by the double line, wants to maintain its armaments at a level equal to or exceeding h' plus a portion l' of US arms. An equilibrium exists where the two security lines intersect, at point E.

[Figure 1 about here]

Figure 1a captures two dynamics associated with an arms race. The first is a *short-run* dynamic and is implicit in the positive slope of each state's security line: if one state increases its level of armament, the other state will react by increasing its own level of armament. The second dynamic is the *long-run* equilibrium which emerges when the two states engage in this short-run action-reaction process. Maintenance of equilibrium in a competitive arms accumulation process requires that a disturbance to this relationship be met by a process of adjustment whereby accumulation patterns return to the equilibrium level (Figure 1b). In the case of a shock to the equilibrium relationship which moves the outcome to point R, both countries find themselves below their minimum security requirements. The

United States and the Soviet Union each increase their armaments to levels indicated by their respective security lines, with the outcome being R!. The process continues through outcome R! until the arms accumulation process returns to equilibrium E. An analogous adjustment occurs when a shock to the equilibrium relationship results in an outcome at S, where both countries have accumulated arms in excess of their security requirements.

Whereas the short-run action-reaction process suggests that increases in one side's armaments have a *direct* impact on the armament level of the other, the notion of reequilibration suggests something different. If a disequilibrium condition is the result of, say, a positive shock in one side's arms level, then the impact on the other side is *indirect* in that the latter is reacting not to the increase in the former's armament level per se, but to the development of a disequilibrium in their relationship. The reacting state would respond similarly if the disequilibrium was brought about by a shock in its own arms level.

Among the most frequently heard criticisms of the Richardson arms race model is that it is mechanistic. While there can be no doubt that the model is simplistic in this sense, haunted by what Etcheson (1989: 9) calls the "ghost of mechanism," most arms-race modelers have not been deterred by such a charge. Indeed, the criticism may be somewhat misplaced. We generally condemn "real world" arms races because one or both sides seem to be reacting to the armaments level of the other side without considering that those armaments may reflect legitimate security concerns, and without considering the security implications of the resulting arms spiral. Statesmen are accused of failing to take into account the strategic intentions of their counterparts, of reducing those intentions to capabilities. In short, we condemn such patterns of arms accumulation for being, well, mechanistic. There may be many things wrong with the Richardson model, but its mechanistic dimension would seem to be one of the things that is right.

Critics have also pointed to the static nature of arms race models. Again, this is a valid observation, but it is misplaced if directed at Richardson's *theoretical* model. The arms race model highlights a dynamic process. A stable equilibrium exists if certain conditions are met (i.e., if $g,h > 0$ and k! l! < 0), and forces are specified (reaction and fatigue) whereby equilibrium is achieved and/or restored.[3] The dynamic process envisioned here is a very simple one, to be sure, but the framework is flexible. If, for example, each side became more sensitive to the arms level of the other, the security lines depicted in Figure 1 would shift upward and outward, resulting in a new equilibrium point at a higher level of armament (see Nicholson 1992: 175-177). In short, the basic Richardson framework can accomodate both disturbances in equilibrium relationships and shifting equilibria. It does not, however, specify precise mechanisms of re-equilibration. It is here where much of the second-generation of arms-race research was focused, particularly in the application of control theory (e.g., Brito and Intriligator 1973; Gillespie and Zinnes 1975; Gillespie et al. 1977).[4]

THE ARMS RACE MODEL IN PRACTICE

For arms-race researchers, empirical analysis of the US-Soviet arms race has been genuinely frustrating. Here is an action-reaction process that would seem to have a great deal of face validity, yet empirical estimation of reaction coefficients has almost always left researchers without scientific evidence for what non-practitioners would consider fairly obvious. One conclusion has been that we have so far failed to get the empirical model specification just right, so subsequent efforts have experimented with alternative functional forms, control variables, empirical indicators, estimation techniques, and the like (see Russett 1983). For instance, in one of more successful efforts to generate empirical evidence for the US-Soviet arms race, Ward (1984), following the suggestion of Taagepera (1979-80), constructs a model which includes a measure of weapons stocks in addition to the traditionally employed budgetary measures.

Another approach to empirical anomalies, particularly for those working within a choice theoretic framework, has been to offer rationalist explanations for otherwise counter-intuitive findings. Williams and McGinnis (1988: 980), for example, suggest that "rational expectations completely reverses our intuitive expectations of the statistical results that would be consistent with the occurrence of an arms race." They argue that although a rival state's previous military expenditures may have affected another state's earlier expectations, the latter's current expectations have been updated on the basis of previously available information. Past expenditures by a rival will not therefore be a good predictor of a state's current expenditures because that information has already been acted upon. More recently, Oren (1996) has proposed a rather ingenious explanation for the *negative* reaction coefficients reported in the empirical arms race literature: to wit, for any given level of hostile behavior, a better armed state will actually appear less threatening than a more lightly armed state, for whom that level of hostile behavior requires more exertion. In short, one cannot help but come away from the literature with the conclusion tha empirical arms-race researchers have not wanted to believe their eyes.

Pitfalls in Empirical Arms-Race Research

Russett (1983: 544) observed that "[t]he greatest number of early works assumed that the change in one state's spending depends on the *level* of its rival's spending in the preceding period." While some may have abandoned this assumption in light of a failure to produce empirical evidence of an arms race, the real effect of using level data is likely to be "false positive" inferences, or type-II errors. Setting aside widespread intuition that the superpower engaged in an arms race during the cold war, if in fact no such action-reaction process operated, ordinary least squares regression using non-stationary time series may well yield evidence of an arms race. For instance, as we will show, the relationship between US and Soviet military *outlays* was not an action-reaction process, but in bivariate regressions using levels, the reaction coefficients are positive and statistically significant.[5] Both outlay series are integrated of order one, and the regression of one non-stationary time series on another is most likely to violate the OLS assumption of serially uncorrelated residuals, with the consequence being unbiased coefficient estimates but deflated standard errors for those estimates. The observed relationship is spurious; we have committed a type-II error.

For both conceptual and statistical reasons, many arms-race researchers have shifted to difference models. As Russett (1983: 544) explains, "a marked increment in a rival's effort may seem threatening even if the level of the rival's effort is low," while a decrement "may constitute a strong signal to call forth comparable restraint from a rival." Difference models are not only conceptually appealing in this light, they may also offer a solution to the statistical inference problem just described. For purposes of empirical estimation, the most basic form is:

$$)X_t \ ' \ (_0 \ \% \ (_1)Y_t \ \% \ <_{1t} \quad [5]$$

$$)Y_t \ ' \ 8_0 \ \% \ 8_1)X_t \ \% \ <_{2t} \quad [6]$$

where $)X_t$ and $)Y_t$ represent changes in rivals' military expenditures from the previous period; $(_0$ and 8_0 are constants, $(_1$ and 8_1 are reaction coefficients, and $<_{1t}$ and $<_{2t}$ are error terms.

Although differencing time series is a common procedure for achieving stationarity, we will argue that it is not always the best approach. If our theoretical model specifies a relationship between periodic *changes* in two rivals' military expenditures (annual or otherwise), then estimating equations [5] and [6] is of course appropriate. But if our theory specifies *levels*, and we estimate [5] and [6] only as a means of imposing stationarity on the data, then we may be led astray. Suppose that X_t and Y_t are random walk series which do in fact derive from an action-reaction process. Regressing first differences of one on first differences of the other is equivalent to regressing one white noise series on another. Clearly, we will find no evidence of an arms race. As a general rule, models estimated with differenced series yield inflated standard errors, thereby undermining our confidence in the stability of estimated coefficients. That is, differencing time series may remove a shared systematic component and increase our chances of committing type-I errors.

AN ERROR-CORRECTION MODEL OF THE ARMS RACE

If two states are engaged in an arms race, then we expect to find their armament levels moving together over time. While point E in Figure 1 captures the outcome at a given point in time, over a span of time we would observe a "moving equilibrium." Time-series analysts have termed this *cointegration*. Such a moving equilibrium in US and Soviet armament levels represents a commonsense understanding of the superpower arms race. Most empirical investigations have started with this commonsense understanding and have sought to reveal the short-run action-reaction dynamic that generates this moving equilibrium in superpower military expenditures (which, of course, is generally considered much too high and a sub-optimal outcome for both sides). That is, the moving equilibrium is assumed while analysis has focused upon whether the slopes of security lines in Figure 1 are indeed positive. Empirical results have been frustrating, as we have mentioned. Regressions have often yielded statistically insignificant reaction coefficients, or, worse, significant ones with negative signs. The typical approach to empirical testing, therefore, has been to estimate the direct effects of one state's military expenditures on those of another.

Even in the absence of the sort of mistaken inferences associated with the use of nonstationary time series, the preoccupation with short-run action-reaction dynamics neglects the important long-run dimension of the arms race. In fact, given that these short-run processes occur in the context of a long-run process of equilibration and reequilibration, why should the same increase in, say, Soviet expenditures elicit the same US reaction regardless of circumstances -- i.e., when there exists an expenditure gap in favor of the Soviet Union, when the gap favors the United States, or when there is rough parity? The quest for stable reaction coefficients would seem to require that we control for long-run dynamics. What is more, even when a proper test of short-run dynamics has been undertaken and the results found wanting, it is premature to conclude that one side's military spending had no effect on the other's. To hypothesize this sort of an action-reaction process is indeed to expect rather much from less-than-omniscient policymakers and their lumbering state bureaucracies (Rajmaira and Ward 1990). However, if the minimum condition for the existence of an arms race is the co-movement in two states' military expenditures over time, then it must be the case that this long-run equilibrium is maintained by one or both. Time-series analysts have termed such behavior *error correction*. Error correction implies indirect effects to the extent that although one state's increased military expenditures may have perturbed the equilibrium relationship, the reaction of the other is to the resulting disequilibrium, to the changed context, and not the to specific action which brought it about.

An error-correction model (ECM) allows for the estimation of re-equilibration parameters, as well as transient dynamics, and so it constitutes a significant advance over existing empirical approaches to arms-race modeling.[6] The methodology is borrowed from econometrics, where it is still evolving. Political scientists have been fairly quick to import the technique; yet there have been relatively few applications in international relations and, as far as we know, no applications to the study of arms races.[7] The notion of equilibrium is basic to an error-correction model of the US-Soviet arms race, and can be defined simply as a state in which there is no inherent tendency toward change. When equilibrium has broken down due to some outside disturbance (or "shock"), the magnitude of the forces of reequilibration will depend on the degree of deviation from the equilibrium state at any given point in time. Two or more phenomena in a long-run equilibrium relationship may experience periods of disequilibrium, but if the equilibrium is stable, and if there are no further shocks, then the relationship will return to its equilibrium state. Operationally, a long-run equilibrium relationship is manifest in the co-movement of time series. An equilibrium relationship, $f(X_t,Y_t)=0$, exists if deviations from equilibrium, $\epsilon_t \equiv f(X_t,Y_t)$, constitute a zero-median stationary process -- i.e., the error term ϵ_t does not exhibit unbounded growth over time.[8]

Cointegration, econometric parlance for co-movement, implies two or more integrated time series. An integrated time series is one for which any given observation represents the accumulation of all past disturbances. The effects of outside shocks never disappear. An integrated series is nonstationary in that there is no mean level to which it ultimately returns and its variance is infinite (only a longitudinal sample has a mean and finite variance). Two or more integrated time series are cointegrated if they share a common stochastic trend -- i.e., they "shadow" one another, never drifting very far apart. Some linear combination of these nonstationary series will produce another, $\epsilon_t = Y_t - (\beta_0 + \beta_1 X_t)$, that is stationary. If there exists a "cointegrating vector," β, which yields a stationary ϵ_t, then we may conclude that the phenomena generating the integrated series are in an equilibrium relationship.

Evidence of cointegration may support the existence of our hypothesized arms race, but we would like to know more. Who reacts to whom? Was the United States responding to Soviet armament, or vice versa? Or was it mutual? Error-correction modeling examines not only the long-term equilibrium relationship (the US-Soviet arms competition more generally), but also short-term effects (the action-reaction process). In its most basic form, an ECM can be represented as follows:

$$\Delta Y_t = \mathrm{T}\Delta X_{t-1} - \alpha(Y_{t-1} - \beta X_{t-1} - C) + \varepsilon_t \qquad [7]$$

Changes in Y are a function of changes in X during the previous period plus the extent to which X and Y were in disequilibrium during the previous period. The equilibrium error is $\{Y_{t-1} - \beta X_{t-1} - C\}$, with C being the spread between the two series in equilibrium, while α is the rate at which a spread of more or less than C is returned to C.

Consider a hypothetical example where policymakers in one state change their military expenditures, Y_t, by exactly the same amount as the previous year's change in their rival's expenditures, X_{t-1}, the only exception being that any deviation from a desired 50 unit advantage in expenditure levels will be corrected at a rate of 80 percent per year. So in this example, $\mathrm{T} = 1$, $\alpha = 0.8$, $\beta = 1$, and C = 50. Figure 2 shows the movement of the two series over 30 years, given a starting value of 10 for X and a 10 unit increase per year. Notice the two shocks and their effects. In year 11, instead of a 10 unit increase, X jumps by 100 units. This creates a disequilibrium -- a spread of -40 instead of +50 -- which policymakers in the responding state seek to correct. But in the process of both matching their rival's increase and adjusting their expenditures to correct the disequilibrium, they "overdo" it, creating a spread of +122 in year 12, a larger advantage than they desire. The process of error correction continues until expenditure levels are reequilibrated. The 50 unit spread has for the most part re-materialized by year 15, although very small adjustments continue to be made. The second shock, to series Y in year 21, involves a 100 unit increase in the responding state's expenditures rather than the 10 unit increase we expect given the rival's spending pattern. In our simple example, the rival is oblivious to changes in the responding state's expenditures, so the shock has no impact on the movement of X. The shock does create a disequilibrium that the responding state itself seeks to correct, and this has largely been accomplished by year 25.

[Figure 2 about here]

In applying the error-correction methodology to the US-Soviet arms race, we do not necessarily impose the same restrictions imposed on equation [7]. An ECM for Y takes this this general form:

$$\Delta Y_t = \mu + \sum_{i=1}^{p} \mathrm{M}_i \Delta Y_{t-i} + \sum_{i=1}^{q} \mathrm{S}_i \Delta X_{t-1} - \alpha\,,_{t-1} + \varepsilon_t \qquad [8]$$

Here, changes in, say, US military expenditures, ΔY_t, are a function not only of past changes in Soviet expenditures, ΔX_t, but also past changes in US expenditures, ΔY_{t-i}. This of course captures Richardson's original notion of "fatigue" or, alternatively, bureaucratic momentum (the difference between the two being the sign of the coefficient).[9] We will want to leave open the possibility that American policymakers do not restrict their consideration to just last year's expenditures (whether US or Soviet), so M_i and S_i represent blocks of coefficients on lags 1 through p and q of ΔY and ΔX, respectively. Finally, we add a constant, μ, and use $,_{t-1}$ to designate the previous period's equilibrium error. As in equation [7], α indicates the rate of reequilibration. It allows for the possibility that in addition to (or in the absence of) the direct effects of changes in Soviet expenditures, those expenditures may have indirect effects in that changes in Soviet expenditures produce deviations in the long-run equilibrium relationship to which US expenditures react. Naturally, we can model changes in Soviet expenditures in an analogous fashion.

It is the equilibrium error term in the ECM that improves its efficiency vis-à-vis alternative methodologies (e.g., vector autoregression). Failure to include such a term may mean the loss of long-run information, especially when data have been differenced to impose stationarity. On the other hand, when series in levels are used to derive estimates, and when two or more of them exhibit co-movement (i.e., share a common stochastic trend), there will be a failure to impose the appropriate cross-equation restrictions and coefficient estimates will have nonstandard distributions (Engle and Granger 1987; Phillips 1991). ECMs use differenced series, but incorporate long-run information via the equilibrium error term. Durr (1992: 187) has remarked that "ECMs provide time-series with what may be a golden mean between the two widely used modeling techniques that focus exclusively on either levels or changes."

DATA ANALYSIS

Military Expenditure Data

Most empirical models of the US-Soviet arms race have operationalized military procurement using defense outlays. Outlays represent the actual spending of funds authorized for military programs and activities. These are the most readily available figures, particularly for countries other than the United States, and when figures are not released by the government, as was the case for the Soviet Union, outlays are easiest to estimate since they are manifest in actual (and presumably observable) defense-related activities. Military expenditure data released by the US Arms Control and Disarmament Agency, which relies heavily on estimates made by the intelligence services, are outlays (see ACDA, annual).

But actual expenditure is only the last phase of a rather prolonged budgetary process, at least in the United States (Ostrom 1977; Majeski 1983). Long before funds are spent, agencies of the Defense Department submit their own budget requests (stage 1). These are adjusted and fit into the overall budget submitted to Congress by the President (stage 2). Congress then authorizes the defense budget, usually after considerable haggling with the President and defense chiefs (stage 3). Only a portion of defense outlays (stage 4) for any given fiscal year are funds authorized by Congress during that same year. Outlays also include obligations incurred during prior fiscal years. Prior obligations are a significant share of the total, since major weapons projects can take quite a few years to fully implement. In fact, only about 75 percent of the funds authorized by Congress for procurement in any given year will have been spent three years hence.[10] For the defense outlays as a whole, perhaps two-thirds represent current-year authorizations; the remainder represents authorizations from previous years.

These observations are relevant for the arms-race hypothesis. If a big chunk of current-year military outlays are beyond the control of current-year policymakers (DoD, the President, or

Congress), we expect that evidence of action-reaction dynamics will be weakest when this indicator is analyzed. The outcomes of each of the three earlier budgetary stages -- service requests, presidential requests, and congressional authorizations -- are more likely to reflect US policymakers' assessments of Soviet armament patterns. It is important to note that action-reaction dynamics may also be relatively weak in Soviet expenditure data, given that our estimates of Soviet expenditures represent outlays. It is not unreasonable to assume that a somewhat analogous budgetary process applied to Soviet defense procurement, even if we do not assume the existence of an autonomous legislative body like the US Congress with power of the purse. We certainly want to hypothesize that Soviet policymakers considered US military spending patterns when formulating their own budget requests, but we need to recognize that evidence for such action-reaction dynamics may be weak when analyzing Soviet outlays.[11] Unfortunately, estimates of Soviet outlays are all we have.

We will test for action-reaction dynamics in four pairs of time series: Soviet military outlays and (1) US service requests, (2) US presidential requests, (3) US congressional authorizations, and (4) US outlays. The series extend from 1948 to 1991.[12]

Data Transformations

Even casual inspection of a plot of US and Soviet military expenditures in levels suggests strongly that the two have moved together over time. The co-movement is most striking when comparing the series in current dollars, but it is clear in the constant-dollar series as well. Indeed, there is no consensus among arms-race modelers as to whether current- or constant-dollar series are most appropriate when testing the arms-race hypothesis. In defending their requests for budgetary increase, US Secretaries of Defense have often referred to constant-dollar trends in US and Soviet military spending, which shows a widening Soviet lead throughout the 1970s. The gap is much less pronounced when viewing the current-dollar series. However, it is generally the case that the US defense budget is proposed, debated, authorized, and spent in current terms, even though some policymakers may take into account constant-dollar trends (Ostrom 1977). Since the issue remains unresolved, we conduct our analyses using both current- and constant-dollar series. We convert to constant dollars using a price deflator constructed specifically for the costs of military goods and services.[13]

In order to capture the long-run equilibrium relationship between US and Soviet military spending, we estimate cointegrating regressions. This requires both that the series be integrated and that they be integrated of the same order, usually $I(1)$. Most of the raw series we examine here are in fact $I(1)$, but Soviet expenditures in current dollars happen to be $I(2)$. A logarithmic transformation of that series eliminates the quadratic trend yielding a series which much more closely resembles the random walk needed for cointegration analysis. Fortunately, a cointegration relationship between series in levels is preserved when the series are log transformed. In fact, it has been suggested that it is easier to detect linear equilibrium relationships among series in logs than series in levels (Granger and Mizon 1993: 192-205). Since we must transform current-dollar Soviet expenditures to obtain a $I(1)$ series, we apply a logarithmic transformation to all other series as well. This entails no loss of information and thus preserves the integrity of our inferences.

Testing for Integration and Cointegration

To restate, an error-correction model treats changes in each series as a function of lagged changes in that series and all other series in the system, plus the previous period's equilibrium error. (The system of interest to us consists only of two phenomena, US and Soviet armament.) It is the inclusion of the prior equilibrium error (ϵ_{t-1} from equation [8]) which distinguishes the ECM approach from vector autoregression (VAR), an increasingly common time-series methodology in international relations research. In essence, ECMs incorporate long-run information that VARs miss (Granger 1988). Engle and Granger (EG) propose a two-step procedure for estimating an ECM (Engle and Granger 1987). First, Ordinary Least Squares is used to estimate a cointegrating regression of the form:

$$Y_t = \beta_0 + \beta_1 X_t + \epsilon_t \qquad [9]$$

This generates a consistent estimate of the cointegrating vector, β, so that we may test whether the equilibrium error series, ϵ_t, is stationary. If so, then we may conclude that *X* and *Y* are cointegrated and that one or both phenomena can be modeled as an error-correction process. In the second stage, the estimated ϵ_{t-1} along with lags of *X* and *Y* are used to estimate the full ECM.[14]

Before estimating the cointegration regression, we want to confirm that our series are *I*(1). To test for the presence of unit roots in our series, we use the Dickey-Fuller (DF) and Augmented Dickey-Fuller (ADF) tests. The Dickey-Fuller test involves regressing a series in first differences on its lagged levels and testing for a statistically significant and negative coefficient, the presence of which implies an autocorrelation parameter less than one and allows us to reject the null hypothesis of a unit root in levels. If the residuals from the DF regression are serially correlated, the coefficient estimate is inefficient and it becomes necessary to include lagged dependent variables (i.e., first-difference terms) as regressors. This is the Augmented Dickey-Fuller test. The top panel of Table 1 shows the results of the DF and ADF tests for each of our log-transformed series in levels.[15] Notice first that the results of the two unit root tests are consistent for all expenditure series in both current and constant dollars. The DF regression for US outlays in current dollars showed evidence of autocorrelation in the residuals, so our inferences are based on the ADF test. With the exception of constant-dollar US outlays, all series show evidence of nonstationarity in levels.

[Table 1 about here]

Since it is possible that nonstationarity derives from a deterministic time trend, we want to distinguish between difference stationary and trend stationary series. The middle panel of Table 1 reports the results of a test suggested by Durlauf and Phillips (1988). The R^2 and Durbin-Watson statistic are from a regression of each series in levels on a time plus a constant. A high R^2 indicates a time trended series (following Nelson and Kang 1983, the rule of thumb for "high" is $R^2 > 0.44$). A high Durbin-Watson (DW) statistic, as confirmed by a Lagrange Multiplier (LM) test, will suggest that the regression is spurious. So although the R^2s in Table 1 are high, the LM statistics indicate that all regressions on time are spurious, allowing us to reject the hypothesis that these series are trend stationary. We are now almost at the point of concluding that our series, with the exception of US outlays in constant dollars, are *I*(1). If that is the case, then the series in first differences will be stationary. The bottom panel of Table 1 shows the DF and ADF results for the differenced series. They are all stationary in first differences.[16] In short, based on the results of unit root and trend stationarity tests, we conclude that all but one of our time series are integrated of order one and that it is sensible to move to the first step of the EG two-step procedure.

[Table 2 about here]

Note first that constant-dollar US and Soviet outlays cannot be cointegrated because the former is stationary in levels. We therefore conduct our tests by estimating cointegrating regressions (equation [9]) for seven pairs of series, treating US expenditures as the endogenous variable. The results are reported in Table 2. It would appear from the cointegrating regression estimates that the relationships between Soviet military expenditures and the various indicators of US defense budgeting are statistically significant.[17] The DF and ADF statistics are for the residuals from the cointegrating regressions and are similar to those used to test the integration of the expenditure series.[18] In three cases, our inferences must be based on the ADF tests since the residuals are autocorrelated. We conclude from the unit root tests that six of the seven pairs of series are cointegrated: *there was a long-run equilibrium relationship between Soviet military outlays and US defense budgeting* (service requests, presidential requests, and congressional authorizations). This is evident in both the current- and constant-dollar series. However, we cannot reject the null hypothesis of no cointegration between Soviet and US outlays.

Constructing and Estimating the Error-Correction Model

Having established the existence of an equilibrium relationship between US and Soviet armament, we have essentially confirmed statistically what scholars and policymakers considered a basic (and fairly obvious) ingredient of the cold war: the superpower arms race. Still, we have not examined the action-reaction process which is presumed to have maintained that equilibrium relationship. Most assume that the arms race was an outcome of a reciprocal dynamic, but partisans may like to attribute to one or another side primary responsibility for the arms race -- by provoking the other with an excessive arms buildup, or by reacting unreasonably to a legitimate buildup. Any conclusion regarding the precise nature of the action-reaction dynamic requires estimation of an error-correction model.

Prior to specifying the final ECM, Engle and Granger (1987) suggest preliminary testing via unrestricted VARs of the form:

$$\Delta Y_t = \mu + \sum_{i=1}^{p} \beta_i \Delta Y_{t-i} + \sum_{i=1}^{p} \gamma_i \Delta X_{t-1} + \eta Y_{t-1} + \xi X_{t-1} + \varepsilon_t \qquad [10]$$

Both and short- and long-run processes are represented here. Estimated coefficients γ_i indicate whether lagged changes in a state's military expenditures have an effect on changes in its rivals expenditures, thus approximating short-term arms-race dynamics. The statistical significance of the γ_i should be examined as a block since collinearity among multiple lags is likely to mislead inferences based on individual coefficient estimates. Estimates of coefficients η and ξ on the lagged equilibrium variables should shed some additional light on the long-run relationship implied by the cointegration tests. For instance, if *Y* represents US expenditures and *X* represents Soviet expenditures, a positive and statistically significant estimate of β reinforces our previous finding regarding the co-movement of the US and Soviet series. A negative and statistically significant estimate of λ constitutes preliminary evidence that US budgeting included an error-correction mechanism.

[Table 3 about here]

Table 3 shows estimates from six VARs. Examination of the residual series from preliminary runs indicated that inclusion of two lags of the difference variables was sufficient to eliminate autocorrelation (i.e., $p = 2$). The estimates for lagged levels of US and Soviet expenditures are consistent across the six VARs (columns 6 and 7). The positive and statistically significant estimates for Soviet outlays in each of the US budgeting equations supports the co-movement hypothesis, which was also supported by the cointegration results. The negative and significant estimates for US spending in these same equations suggests that US policymakers tended to correct for deviations from the equilibrium relationship, and therefore that the US budgeting series should be represented as error-correction processes. Notice that neither US nor Soviet expenditures in levels are found to be statistically significant in any of the Soviet outlays equations. Linear combinations of the equilibrium variables had no impact on changes in Soviet outlays. Here we have our first bit of evidence that the superpower arms race was one-sided.[19]

In moving on to examine short-run dynamics, we can now set aside the equations for Soviet outlays, which is found to be (weakly) exogenous to US service requests, presidential requests, and congressional authorizations in both current and constant dollars. We focus our attention on the US spending equations. Table 4 reports the results of tests of joint significance for the two lags of US and Soviet spending in first differences. The second column of F values provides evidence that last year's changes in US service requests and congressional authorization "Granger cause" current year changes. The joint significance tests do not suggest that lagged changes in Soviet expenditures have had a direct impact on changes in US budgeting (with the possible exception of US presidential requests, where the F value is statistically significant at the 0.10 level). These results seem to indicate that Soviet outlays did not generally have a direct impact on US budgeting in the short-run, but rather that the transient dynamics were dominated by bureaucratic momentum.[20]

[Table 4 about here]

The VAR results inform the specification of our error-correction models. First, to reiterate, we need not specify an ECM for Soviet outlays since the VARs yield evidence of error-correction dynamics in US defense budgeting only. Second, we are now able to place several restrictions of the general ECM model (equation [8]). The models are specified as follows:

Service Requests	$\Delta Y_{1t} = \mu_1 + \phi_{11}\Delta Y_{1t-1} + \phi_{12}\Delta Y_{1t-2} - \alpha_1 z_{1t-1} + \varepsilon_{1t}$	[11]
Presidential Requests (current)	$\Delta Y_{2t} = \mu_2 - \alpha_2 z_{2t-1} + \varepsilon_{2t}$	[12]
Presidential Requests (constant)	$\Delta Y_{3t} = \mu_3 + \theta_{31}\Delta X_{3t-1} + \theta_{32}\Delta X_{3t-2} - \alpha_3 z_{3t-1} + \varepsilon_{3t}$	[13]
Congressional Authorizations	$\Delta Y_{4t} = \mu_4 + \phi_{41}\Delta Y_{4t-1} + \phi_{42}\Delta Y_{4t-2} - \alpha_4 z_{4t-1} + \varepsilon_{4t}$	[14]

The equations for service requests and congressional authorizations (in current and constant dollars) include lagged endogenous variables to capture the effects of bureaucratic momentum. Only presidential requests in constant dollars are modeled as a function of lagged changes in Soviet outlays. All six US series are modeled using a constant and, of course, a lagged equilibrium error term, $,_{nt\text{-}1}$. Whereas the VARs for US budgeting suggest the existence of an error-correction process, given the negatively signed and statistically significant coefficient estimates on lagged US budgeting in levels, our final judgment regarding the nature of reequilibration rests on an examination of the estimated " coefficients from equations [11] through [14]. These estimates indicate the rate at which US policymakers correct for deviations in the equilibrium relationship between US and Soviet armament; therefore, 0 < " < 1.

[Table 5 about here]

Estimates from the six ECMs are reported in Table 5. We see from the last column of estimates that *each of the six US defense budgeting series exhibit an error-correction mechanism*. The " coefficient estimates are negatively signed and statistically significant, as predicted. Reequilibration appears to have occurred most rapidly in the case of US service requests, with deviations in the equilibrium relationship corrected at the rate of 68 and 75 percent per year (current- and constant-dollar series, respectively). Error correction in presidential requests and congressional authorization occurred at rates from 38 to 65 percent per year. The somewhat faster correction rates for service requests may reflect fewer competing interests involved at that stage of the defense budgeting process. The President and then Congress must situate defense spending alongside a large array of (mostly) domestic programs, so we should not be surprised to find that presidential requests and congressional authorizations are not quite so sensitive to deviations from the long-run equilibrium.

As far as transient effects, notice that of the four series exhibiting short-run momentum and/or fatigue effects in the VAR analysis, only congressional authorizations in current dollars exhibit the same when modeled as an error-correction process. Taken together, lags of congressional authorizations in first differences have a statistically significant impact on current-year authorizations and, judging from the negatively signed coefficient estimates, the short-run effect being captured here is fatigue and not momentum. As a representative body, the US Congress is presumably not driven by bureaucratic momentum, at least on the matter of military spending. Finally, recall that the VAR results suggest that only the ECM specification for US presidential requests in constant dollars should include lagged changes in Soviet outlays. The F test does indeed indicate that the short-run impact of Soviet outlays was statistically significant, with coefficient estimates being positive as the arms-race hypothesis predicts. Military service requests probably reflect more a preoccupation with respective bailiwicks and less a desire to match Soviet spending increases in the aggregate.[21] It is the Secretary of Defense who is more likely to be sensitive to changes in aggregate Soviet spending, and his influence is most potent when the President is preparing budget requests for Congress.

CONCLUSIONS

The superpower arms race consisted of both long-run and short-run dynamics. Six pairs of time series -- Soviet military outlays and US service requests, presidential requests, and congressional authorizations in both current and constant dollars -- exhibited evidence of cointegration, which suggests that US and Soviet armament efforts maintained a long-run equilibrium relationship. That is the pattern we usually associate with an arms race. But the relationship was not symmetrical, for it was the US policymakers and not the Soviets who adjusted their defense budgets in the face of deviations from equilibrium. Each of the six US budgeting series, but not Soviet outlays, manifest an error-correction mechanism. The superpower arms race appears from this evidence to have been one sided: the Soviets acted, the Americans reacted. Relative to these long-run dynamics, evidence for US reactivity in the short-run is sparse. Only presidential requests in constant dollars were a direct response to previous changes in Soviet outlays. Overall, then, the impact of Soviet outlays on US defense budgeting was mostly indirect -- i.e., to the extent that Soviet policy caused deviations in the equilibrium relationship, which US policymakers sought to correct.

Asymmetry in US and Soviet behavior is a finding which has surfaced in other studies, including several employing time-series methods. In their examination of reciprocity, Ward and Rajmaira (1992: 354) conclude that "U.S. conflict behavior during [the 1948-1988] period was a function of the conflict it received from the Soviet Union," while "the Soviet Union did not pay much attention to the level of conflict it received from the United States." Two of three datasets analyzed by Goldstein and Freeman (1990: 74-76) suggest that although the United States reciprocated Soviet behavior (conflictual and cooperative), the Soviet Union did not react similarly. Action-reaction dynamics have also been examined in superpower arms transfers to the Third World. Kinsella (1995) reports evidence that US arms-transfer policy was reactive to Soviet transfers to South Asia, the Middle East, and the Persian Gulf, but that only in the latter region was reactivity mutual (see also Kinsella 1994). So the quantitative empirical literature on US-Soviet relations would appear to suggest that the one-sided nature of the superpower arms race is not unprecedented when considered in a more general context.

We do not wish to rush to that conclusion, however. In fact, our test of the arms-race hypothesis is biased against finding Soviet reaction, or error-correcting behavior, for reasons we alluded to above. The only Soviet military expenditure data available represent estimates of Soviet defense outlays. We need only assume a rough similarity between the US and Soviet defense budgeting process to hypothesize that Soviet outlays were weakly correlated with US defense budgeting. In the Soviet Union, expenditures approved for procurement were no doubt spread over the course of a few years or more, as they were (and are) in the United States. Therefore, actual outlays in any given year were probably not as sensitive to external forces like changes in US defense spending, or to shocks to the long-run equilibrium in the two states' armament levels. We were not terribly surprised to find that US and Soviet defense outlays did not maintain a long-run equilibrium relationship (i.e., were not cointegrated), given the lesser degree of discretion exercised by US policymakers on the matter of current-year outlays versus requests and authorizations. We can reasonably assume the same for Soviet outlays. Without data on Soviet defense *budgeting*, as opposed to (estimates of) actual expenditures, we are ill-equipped to fully examine arms-race dynamics.

It is somewhat ungratifying to conclude yet another study of the US-Soviet arms race without having presented definitive quantitative evidence as to its precise nature. But the error-correction methodology has allowed us to zero in on what is perhaps the last impediment to a balanced and parsimonious empirical model of the arms race: Soviet defense budgeting data. We have been able to model the long-run as well as short-run dynamics in the US and Soviet expenditure data we do possess, and have thereby undertaken a more complete analysis than earlier studies which restrict themselves to one or the other (usually the latter). At the same time, the model is quite parsimonious in the sense that arms-race dynamics have been captured in expenditure data alone, without recourse to more complex indicators of US and Soviet defense efforts or various other control variables. Action-reaction dynamics are right there in the US defense budgeting data; the error-correction methodology has brought them out. We suspect that action-reaction dynamics may also be there in the Soviet budgeting data, but that must remain a working hypothesis.

ENDNOTES

1. While this basic action-reaction model has received the most analytical and empirical scrutiny in the literature, Richardson also proposed some variants of this model. In his "rivalry model," the defense coefficients k and l apply not to the absolute arms level of the opponent, but rather to the armaments gap obtaining between the two sides. His "submission model" includes an additional term for the tendency of one side to "back off" in the face of a widening armaments gap in favor of its opponent. See Etcheson (1989: 32-35).

2. A somewhat more elaborate discussion of these matters can be found in Intriligator and Brito (1989).

3. The model is consistent with other patterns as well, in which equilibria are either unstable or nonexistent. If, for example, both sides recognized that the other had security concerns unrelated to their own rivalry, the grievance terms g and h would in essence become negative. Even when disarmed, each side tolerates some amount of arms possessed by the other. In Figure 1, it is as if the two security lines were switched. Point E is still an equilibrium, but it is unstable: a disturbance in the relationship to, say, point S would lead to an explosive arms spiral. See Nicholson (1992: 171-175).

4. As a general model of the arms accumulation process, the Richardson framework is of course far from adequate. The fatigue component notwithstanding, there is no real attempt to represent the internal bureaucratic dynamics at work in determining a state's level of military spending (see especially Ostrom 1977, 1978 and Majeski 1983). Although this is perhaps the model's most serious limitation, other characteristics of the external action-reaction dimension of arms accumulation are not explicit in the model -- e.g., time lags between reactive budget allocations, nonlinear relationships, etc. On the other hand, these would seem to be empirical issues and best not incorporated into a theoretical model until it has been confirmed that they characterize most "real world" arms races (and in what form). The empirical study of arms races has not yet yielded confident observations on these matters.

5. The results of statistical exercises, such as this one, alluded to but not reported in this paper may be found in Chung (1996).

6. The error-correction methodology is detailed in Engle and Granger (1987). For a review of the technique in the context of political research, see the symposium in *Political Analysis*, volume 4, especially articles by Ostrom and Smith (1992) and Durr (1992).

7. One partial exception is a study by Kollias (1996) examining Greek military expenditures in the context of the Greek-Turkish rivalry. While the analysis does show evidence of an error-correction mechanism in Greek military expenditures, the equilibrium relationship being maintained involves more that Turkish military expenditures (i.e., Greek GDP). Kollias (1996) does not examine Turkish expenditures for error-correction behavior. An error-correction model is also employed by Rajmaira and Ward (1990; see also Ward and Rajmaira 1992) to investigate the nature of reciprocal behavior between the United States, the Soviet Union, and China during the cold war, but the focus here is on

cooperative and conflictual behavior generally and not defense spending. The arms-race literature which incorporates notion of feedback and control shares a basic conceptual element with the error-correction approach. One such empirical study is Desai and Blake (1981).

8. Note that for any two stationary time series, $,_t$ is a stationary process and thus the phenomena will appear to be in equilibrium, even if they are in fact unrelated. Under these circumstances, $,_t$ provides no real information about the relationship. For nonstationary time series, deviations of the form $,_t=Y_t$-bX_t constitute a nonstationary process whenever the observed relationship, b, does not equal the true relationship, $\$$. Only the true relationship yields a stationary $,_t$. See Granger and Mizon (1993: 2-5).

9. Including lags of the dependent variable as regressors also serves to minimize the degree of serial correlation in the error term, which would otherwise attenuate the standard errors of the estimates and increase the likelihood of erroneous inferences. Use of lagged dependent variables is sometimes frowned upon as atheoretical, but in the case of arms-race modeling there is in fact a theoretical justification for this sort of specification.

10. The Congressional Budget Office has reported that, for the United States, "on average, $1 authorized for defense procurement produces only about 12 cents of actual outlays in the first year. Outlays grow to 37 cents in the second year, 30 cents in the third year, and then tail off" (quoted in Chan 1985: 427).

11. It would be reasonable to suggest, however, that the lack of an autonomous legislative body meant that service requests were more likely to be reflected in Soviet military budgets than was the case in the United States. Still, although that implies a higher correlation between service requests and outlays, it does not change the fact that a significant portion of Soviet current-year outlays are likely to represent funding authorizations made in previous years, as in the United States.

12. Soviet outlays from 1967 are assembled from issues of ACDA's *World Military Expenditures and Arms Transfers* (U.S. Arms Control and Disarmament Agency, annual); estimates prior to that are reported in the DoD's *Annual Report to the President of the United States*, fiscal year 1978 (U.S. Department of Defense 1977: 12-26). US service requests are compiled from various issues of the DoD's *Annual Report*; both presidential requests and congressional authorizations from *Congressional Quarterly Almanac*, 1948-1991; and US outlays from *Historical Tables: Budget of the United States Government*, fiscal year 1994 (Executive Office of the President, 1993).

13. The military price deflator is released by the Bureau of Economic Analysis of the Department of Commerce. Despite the availability of this deflator, it is somewhat surprising to find researchers using such alternatives as the consumer or producer price index. Although some of the increase in the costs of military goods (and services) is certainly due to increases in unit costs (i.e., inflation), increases in weapons costs also reflect quality improvements. We therefore want a deflator which attempts to exclude quality improvements from real price increases. The military price deflator does in fact show a flatter trend over time than, say, the CPI. Interestingly, when Defense Secretary Caspar Weinberger

sought to emphasize to Congress the growing gap between Soviet and American military spending, he chose to report figures converted to constant dollars using the CPI. That suggested a growth pattern in US expenditures somewhat flatter than what would have been the case if the military price deflator had been used. Obviously, the CPI better served the Secretary's purposes. See the Defense Department's *Annual Report to the Congress of the United State*, fiscal year 1984 (DoD 1983: 20).

Because our data for Soviet expenditures are estimates made by US intelligence services rather than figures released by the Soviet government, we apply the military price deflator here as well. We should also acknowledge that there is some debate surrounding the procedures used to estimate Soviet military expenditures, including cost increases (see Prados 1982; Noren 1995). Nonetheless, these are the only data that are available over the entire time period of our study. The Stockholm International Peace Research Institute stopped trying to estimate Soviet military spending from 1986 (SIPRI 1987).

14. Some analysts have questioned the necessity of the first step of the EG two-step procedure. Beck (1992) argues that the sorts of series examined by political scientists (as opposed to economists) are not likely to be random walks, or "long memoried," and that it is difficult to distinguish between stationary and integrated processes in any case. Series which appear to be stationary by standard unit root tests may nonetheless be long memoried (with roots close to but less than one). Therefore, according to Beck, we should not be deterred from estimating an ECM (second step) simply because our series do not pass the cointegration test. If they are long memoried, an ECM is appropriate. In fact, Beck suggests proceeding directly to the second step of the EG two-step procedure. Instead of including the equilibrium error series from the cointegrating regression in the ECM, we may simply examine the estimated coefficient from the once lagged dependent series in levels. If it is negative and statistically significant, we accept the error-correction hypothesis (see also Durr 1992). Since Beck is largely concerned with suggesting an approach when series are long memoried but stationary, this simpler single-equation method is sensible. But our series are nonstationary (see below), so this approach would involve regression on nonstationary series (US and Soviet military expenditures in levels), thereby violating the stationarity assumption upon which inferences are based. The EG two-step method avoids this problem since the error series from the cointegrating regression is stationary. For this reason, we stick with the EG two-step procedure.

15. Because the DF and ADF tests may entail regression on a nonstationary series (i.e., lagged levels), the limiting normal distribution for Student's t does not apply. Adjusted critical values, as reported by Charemza and Deadman (1992: 325-328), are as follows: 0.01 level, -3.17 and -3.18 for DF and ADF, respectively; 0.05 level, -2.16 and -2.17.

16. Both current- and constant-dollar Soviet outlays are stationary in first differences only according to the DF test. The ADF test indicates stationarity, but only at the 0.05 level.

17. Again, since the Soviet outlay series constitutes a nonstationary regressor, the standard Student t distribution does not apply, and we should exercise caution when interpreting individual coefficients.

18. The only difference is that in adopting critical values, we must take into account the fact that the cointegrating vector is not known and must be estimated. In this case, the standard unit root tests would be biased in favor of rejecting the null hypothesis of nonstationarity. We therefore adopt critical values which allow for the estimation of the cointegration coefficient, but we relax our confidence criterion to the 0.05 level. That seems to us a prudent approach until more research has been done on the appropriate limiting distributions for DF and ADF tests. Critical values specifically computed for cointegration tests are: 0.01 level, -4.11 and -4.12 for DF and ADF, respectively; 0.05 level, -3.38 for both DF and ADF. See Charemza and Deadman (1992: 325-328).

19. In advocating single-equation ECMs, Beck (1992: 242) points out that because there is often "a clear division of variables into caused and causal" in political science (as opposed to economics), the EG two-step approach allows for "theoretically implausible reequilibrating mechanisms." In that context, a single-equation ECM is both more consistent with theory and easier to estimate. In the case of arms-race research, and research on international reciprocity generally, the behavior being modeled is hypothesized to be both cause and effect, so the symmetric treatment of variables in the EG two-step procedure is appropriate. So, while our empirical results suggest a certain causal asymmetry in US and Soviet armament, the theory informing our analysis does not.

20. Joint significance tests for lagged levels of US spending in the Soviet outlays equations (not shown) indicate that no indicator of US defense budgeting Granger causes Soviet outlays. This, combined with the absence of any transient effects of US budgeting on Soviet outlays, makes Soviet outlays strongly exogenous to US budgeting. This strong exogeneity means that standard distribution theory applies to the ECM estimates from the EG two-step procedure (see Ostrom and Smith 1992: 147-148; Phillips 1991).

21. The F ratio for lagged changes in service requests in the constant-dollar service-request equation is statistically significant only at the 0.10 level. The coefficient estimates are both positive, lending support to the bureaucratic momentum interpretation (and, of course, applying a relaxed criterion for statistical significance).

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Table 1: Tests for Nonstationarity in US and Soviet Military Expenditures, 1948-1991

Test for Unit Root in Levels

	current dollars		constant dollars	
	DF	ADF	DF	ADF
US Service Req.	-1.26	-1.52	-2.38	-2.58
US President Req.	-0.92	-1.72	-2.50	-2.24
US Congress Auth.	-1.15	-1.36	-2.32	-2.34
US Outlays	-2.78†	-2.46	-4.21*	-4.17*
Soviet Outlays	-1.49	-1.00	-2.63†	-1.73

Test for Trend Stationarity

	current dollars			constant dollars		
	R^2	DW	LM	R^2	DW	LM
US Service Req.	0.89	0.78	15.4*	0.77	0.81	13.0*
US President Req.	0.92	0.52	21.7*	0.84	0.57	17.7*
US Congress Auth.	0.92	0.98	10.3*	0.83	0.98	8.9*
US Outlays	0.89	0.29	24.9*	0.76	0.27	22.6*
Soviet Outlays	0.97	0.22	34.9*	0.95	0.29	29.5*

Test for Unit Root in First Differences

	current dollars		constant dollars	
	DF	ADF	DF	ADF
US Service Req.	-6.55*	-4.86*	-6.31*	-4.71*
US President Req.	-5.60*	-4.87*	-4.86*	-4.93*
US Congress Auth.	-7.97*†	-5.99*	-7.61*	-5.81*
US Outlays	-3.88*	-5.09*	-3.78*	-3.48*
Soviet Outlays	-4.15*	-2.72	-4.04*	-2.49

Note: All series are natural logarithms. DF and ADF test statistics designated with an asteriks (*) indicate rejection of the null hypothesis of a unit root at the 0.01 level. DF statistics designated with a cross (†) indicate that regression residuals showed evidence of autocorrelation. ADF regressions include one lag of the dependent variable.   LM statistics designated with an asteriks (*) indicate rejection of the null hypothesis of trend stationarity at the 0.01 level.

Table 2: Cointegration Tests for US and Soviet Military Expenditures, 1948-1991

	Estimates		*Goodness of Fit*		*Cointegration Tests*	
	Constant	Sov.Outlays	R^2	D-W	DF	ADF
current dollars						
US Service Req.	1.18 (2.25)	0.90 (19.47)	0.90	0.86	-3.55*	-3.69*
US President Req.	0.34 (0.74)	0.96 (23.96)	0.93	0.59	-3.05†	-3.57*
US Congress Auth.	0.16 (0.31)	0.97 (21.39)	0.91	0.93	-3.74*	-3.34
US Outlays	1.28 (2.30)	0.88 (17.96)	0.88	0.22	-2.97†	-2.84
constant dollars						
US Service Req.	1.75 (2.25)	0.86 (12.78)	0.79	0.86	-3.59*	-3.80*
US President Req.	0.57 (0.85)	0.94 (16.24)	0.86	0.59	-3.09†	-3.63*
US Congress Auth.	0.32 (0.42)	0.96 (14.59)	0.83	0.93	-3.76*	-3.36

Note: Numbers in parentheses are t-ratios. DF and ADF test statistics designated with an asteriks (*) indicate rejection of the null hypothesis of a unit root at the 0.05 level. DF statistics designated with a cross (†) indicate that regression residuals showed evidence of autocorrelation. ADF regressions include one lag of the first-difference term.

Table 3: Vector Autoregression Results for US and Soviet Military Expenditures, 1948-1991

	Estimates							*Goodness of Fit*	
	Constant	) Sov.(-1)	) Sov.(-2)	) US(-1)	) US(-2)	Sov.(-1)	US(-1)	R^2	D-W
current dollars									
US Service Req.	1.62	0.58	-0.32	-0.52	0.18	0.68	-0.82	0.51	2.05
	(2.78)	(1.07)	(-0.54)	(-3.12)	(1.16)	(3.72)	(-3.88)		
Soviet Outlays	0.17	0.20	0.34	-0.03	-0.09	0.01	-0.02	0.57	1.67
	(0.89)	(1.17)	(1.80)	(-0.48)	(-1.74)	(0.17)	(-0.33)		
US President Req.	0.68	0.69	0.02	-0.18	-0.09	0.48	-0.54	0.61	2.06
	(2.04)	(1.94)	(0.07)	(-1.13)	(-0.66)	(3.27)	(-3.42)		
Soviet Outlays	0.19	0.17	0.34	-0.07	-0.06	0.05	-0.06	0.54	1.78
	(1.19)	(0.95)	(1.68)	(-0.84)	(-0.84)	(0.63)	(-0.78)		
US Congress Auth.	0.74	0.39	0.15	-0.68	-0.07	0.64	-0.70	0.47	2.00
	(1.46)	(0.70)	(0.22)	(-4.07)	(-0.45)	(2.96)	(-3.09)		
Soviet Outlays	0.14	0.19	0.43	-0.07	-0.08	0.00	-0.01	0.57	1.72
	(0.94)	(1.12)	(2.22)	(-1.50)	(-1.66)	(0.01)	(-0.16)		
constant dollars									
US Service Req.	2.48	0.76	-0.15	-0.50	0.15	0.56	-0.77	0.39	2.09
	(2.70)	(1.41)	(-0.27)	(-3.04)	(1.01)	(3.57)	(-3.88)		
Soviet Outlays	0.33	0.19	0.34	-0.02	-0.09	-0.02	-0.02	0.22	1.66
	(1.14)	(1.10)	(1.88)	(-0.35)	(-1.93)	(-0.17)	(-0.30)		
US President Req.	1.18	0.80	0.09	-0.16	-0.11	0.42	-0.52	0.50	2.08
	(2.21)	(2.23)	(0.22)	(-1.06)	(-0.82)	(3.25)	(-3.54)		
Soviet Outlays	0.17	0.15	0.34	-0.05	-0.07	0.02	-0.05	0.16	1.75
	(0.89)	(0.87)	(1.79)	(-0.67)	(-1.01)	(0.33)	(-0.73)		
US Congress Auth.	1.30	0.55	0.32	-0.65	-0.09	0.53	-0.64	0.33	2.01
	(1.58)	(0.95)	(0.49)	(-3.93)	(-0.56)	(2.72)	(-2.97)		
Soviet Outlays	0.30	0.17	0.42	-0.07	-0.08	-0.01	-0.01	0.21	1.69
	(1.27)	(1.01)	(2.31)	(-1.46)	(-1.81)	(-0.24)	(-0.18)		

Note: Numbers in parentheses are t-ratios. All endogenous variables are first differences.

Table 4: Joint Signfiicance Test for Difference Variables in US Budgeting Equations

	F values	
	) Sov.	) US
current dollars		
US Service Req.	0.72	4.94*
US President Req.	1.89	1.33
US Congress Auth.	0.27	8.51**
constant dollars		
US Service Req.	1.02	4.67*
US President Req.	2.51	1.39
US Congress Auth.	0.59	8.01**

Note: F values disignated with two asteriks (**) are significant at the 0.01 level; one asteriks (*) indicates significance at the 0.05 level. Statistics are computed for estimates on lags 1 and 2 of the difference variables reported in Table 3.

Table 5: Error-Correction Models for US Defense Budgeting, 1948-1991

	Estimates						*Goodness of Fit*		*Joint Significance*	
	Constant	) US.(-1)	) US.(-2)	) Sov.(-1)	) Sov.(-2)	Eq. Error(-1)	R^2	D-W	F,) US	F,) Sov.
current dollars										
US Service Req.	0.05	0.21	0.14			-0.68	0.42	1.98	1.68	
	(1.35)	(1.62)	(1.08)			(-5.16)				
US President Req.	0.07					-0.47	0.34	1.41		
	(3.11)					(-4.65)				
US Congress Auth.	0.11	-0.32	-0.25			-0.42	0.24	1.32	3.20*	
	(2.62)	(-2.15)	(-1.77)			(-2.83)				
constant dollars										
US Service Req.	0.04	0.27	0.16			-0.75	0.44	1.96	2.42	
	(1.12)	(1.99)	(1.21)			(-5.32)				
US President Req.	-0.00			0.76	0.52	-0.38	0.45	1.55		4.28*
	(-0.04)			(2.02)	(1.49)	(-3.30)				
US Congress Auth.	0.06	0.04	-0.05			-0.65	0.38	1.94	0.13	
	(1.65)	(0.26)	(-0.34)			(-4.31)				

Note: Numbers in parentheses are t-ratios. F values designated with an asteriks (*) are significant at the 0.05 level. All endogenous variables are first differences.