Accounting for National-Level Public Support for European Integration: Comparing a New Method to Existing Methods for Time-Series Cross-Section Data

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Abstract
This paper addresses two topics: public support for European integration and the statistical analysis of pooled time-series cross-section (TSCS) data. With respect to the first topic, previous research has largely found that national-level public opinion on European integration varies little with national economic conditions and is largely attributable to nation-specific factors. We develop and evaluate empirically a political-economy model in which public support varies consistently with national differences in the economic and political benefits of European Union (EU) membership. Our political-economy model explains over 90% of the cross-national and cross-temporal variation in support without relying extensively on national dummy variables.

We test our model using TSCS data, for which previous research has proposed several methods of panel-robust estimation. We contribute to this literature by adapting the Newey-West heteroscedasticity-autocorrelation consistent covariance estimator to the TSCS context. The principal advantage of this method over others is that it does not require researchers to specify and impose a time-series process in order to derive robust coefficient standard errors. The paper’s statistical analysis compares four alternative panel-robust methods for TSCS data: two versions of the Beck-Katz method, Greene’s method for weighted least squares, and our proposed panel adaptation of the Newey-West estimator. We discuss the relative merits of each method in the specific context of our substantive topic, as well as under different applied circumstances more generally.

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Public opinion, through its impact on mass political behavior, shapes European integration and European Union (EU) policy-making. The influence of mass attitudes is most apparent in the aggregate at the national level. Traditionally, the link between public opinion and EU policy-making has been indirect, through national representative channels. Two of the most important decision-making institutions for the EU—the Council of Ministers and the European Council—consist of representatives of the national governments. These policy-makers are therefore sensitive to the preferences of their national electorates. Research on EU budgetary politics, for example, indicates that national public support for integration influences bargaining among member-states over the distribution of EU resources (Carrubba 1997).

The process of integration is also shaped by national-level public opinion. All treaty revisions require ratification by national parliaments, which are accountable to the national publics. Furthermore, where national referendums on integration are required, national publics exercise direct control over the integration process. For example, the Danish referendum on the Maastricht Treaty on European Union effectively halted the integration process. Several national referendums will also be held on the adoption of Economic and Monetary Union. Consequently, explaining differences in national public support for integration is important for understanding the process of European integration and EU policy-making.

Previous research on national variation in support for integration offers some limited systematic explanations.\(^1\) Consistent with economic voting models employed in comparative

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\(^1\) We refer strictly to the aggregate-level research. There is also a growing literature focusing on individual-level variation in support for integration (e.g., Gabel 1998a; Anderson and Reichert 1996; Gabel and Whitten 1996; Gabel and Palmer 1995).
politics, Eichenberg and Dalton (1993) and Anderson and Kaltenthaler (1996) have shown that national public support for integration varies with national economic conditions. In addition, Vaubel (1994) and Eichenberg and Dalton (1993) found public support related to commercial interaction among the EU member-states. These studies are limited, however, by their reliance on nation-specific explanations to account for much of the cross-national variation. While the publics of the EU member-states are clearly marked by distinct national traditions, we would like to account for these distinctions with general theoretical explanations. Put another way, in our explanations of variation in national public support we would like to replace the names of countries with the names of theoretically meaningful variables (Preworski and Tuene 1970).

In this paper, we develop and test a theoretical explanation for national public support for European integration. Specifically, we contend that cross-national variation in public support reflects national differences in the economic and political benefits of EU membership. In the first section of the paper, we discuss previous research on public opinion toward European integration. We then develop our theoretical explanation in two steps. First, we briefly discuss an individual-level model of consistent opinion formation that could produce the systematic national differences in public support that are revealed by our statistical analysis. Second, we develop a political-economy model that specifies several testable hypotheses that account for national differences in public support for integration.

In the remainder of the paper, we investigate the validity of our hypothesized relationships

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2 An empirical investigation of this individual-level process in an EU context would be a valuable contribution to the public opinion literature, but is beyond the scope of the present paper. Even so, the national-level results presented here have some implications for understanding the nature of attitudes
by analyzing pooled time-series cross-section (TSCS) data composed of aggregated
Eurobarometer survey responses and national-level economic and political indicators from 1975-
1992. The pooled design has been the focus of considerable research because it poses
methodological problems that may undermine statistical inference (Beck and Katz 1995; Greene
1997: 648-69). Due to recent innovations, though, several alternative methods have been
proposed that employ panel-robust estimation of the coefficient standard errors. We contribute to
this literature by adapting the Newey-West heteroscedasticity-autocorrelation consistent
covariance estimator to the TSCS context. The principal advantage of this method over others is
that it does not require researchers to specify and impose a time-series process in order to derive
robust standard errors. To our knowledge, we are the first to propose the application of a panel-
adapted Newey-West covariance estimator.

In our statistical analysis, we apply four alternative methods of panel-robust estimation in
order to compare the empirical evidence for our hypotheses across estimation methods. The four
methods are: (1) ordinary least squares (OLS) coefficients derived with the inclusion of a lagged
dependent variable and evaluated with Beck-Katz standard errors; (2) OLS coefficients evaluated
with panel-adapted Newey-West standard errors; (3) maximum likelihood (ML) coefficients
derived with an AR(1) error process and homoscedasticity imposed, and evaluated with Beck-
toward integration, to the extent that an ecological fallacy in inference does not exist.
3 The Eurobarometer data were originally collected by Jacques-Rene Rabier, Helene Riffault, and Ronald
Inglehart, and made available by the Inter-University Consortium for Political and Social Research.
Neither the collectors of the Eurobarometer data nor the Consortium bear any responsibility for the
interpretations presented here. The economic data were compiled by the authors using two Organization
for Economic Cooperation and Development (OECD) sources: *Foreign Trade by Commodities* and *Main
Economic Indicators*. 
Katz standard errors; and (4) ML coefficients derived with an AR(1) error process and allowing for groupwise heteroscedasticity, and evaluated with Greene’s panel-robust standard errors for weighted least squares. The first and third are different applications of a method proposed by Beck and Katz (1995). Greene (1997) proposed the fourth; and the second is proposed and developed in the present paper. Even though all four methods employ panel-robust standard errors, they assume different structures for the latent causal process. We discuss these structural differences and provide some guidance on how to choose among the methods, in practice, based on the applied circumstances.

The results of our statistical analysis show that a nation's political and economic interests in EU membership determine the strength of support among its citizens. More specifically, our statistical model demonstrates that a national public's support for integration is strongly related to its international trade interests and internal and external stability concerns. Our statistical model explains over 90% of the cross-national and cross-temporal variation in public support from 1975-1992 without relying extensively on national dummy variables. In other words, our model not only fits the observations accurately, it also explains public support for integration with specified comparative relationships that are generalizable across countries rather than with unspecified national effects that are idiosyncratic to particular countries.

I. Public Opinion on European Integration

Past research on public opinion toward European integration has demonstrated several important empirical regularities, but has offered only limited theoretical explanations for these
findings. Probably the most consistent finding of survey research is that EU citizens demonstrate little interest in or knowledge of European integration or EU policy (Slater 1983; Sinnott and Niedermayer 1996). This public apathy suggests that public opinion regarding integration is superficial and transitory. Consequently, variation in national public support for integration may reflect temporary and unpredictable swings of emotions rather than systematic and predictable reactions to integration or its consequences.

Yet, in spite of this apathy, national public support for integration appears to vary in systematic ways. There are discernable trends in public support both across time and across nations (Inglehart and Rabier 1978; Mathew 1980; Inglehart, et al. 1991; Eichenberg and Dalton 1993; Vaubel 1994; Anderson and Kaltenthaler 1996). In particular, many studies find that national public support is stronger in the original members or, put another way, increases with length of membership.

Several studies have attempted to explain these trends in support. Eichenberg and Dalton (1993) tested an economic voting model positing that public support for integration varies with national economic conditions, national dependence on EU trade, East-West conflict, and national return from the EU budget. This model offered some limited explanation for variation in public opinion from 1973-1988 among the first nine EU members. As predicted, decreasing inflation and increasing national dependence on EU trade were positively related to support and were statistically significant. However, the substantive significance of these variables was small. The variables with the strongest substantive impact on national variation in support were national
dummy variables.\textsuperscript{4} Eichenberg and Dalton (1993) attributed these nation-specific effects to different “national traditions” regarding European integration. They claimed, for example, that the relatively high support associated with the France dummy variable reflected growing French support of Europe in the 1980s due to concerns about international economic pressures and the European security environment (Eichenberg and Dalton 1993: 525).

Anderson and Kaltenthaler (1996) re-examined these same hypotheses with data including all of the EU member-states from 1975-1993. They found that improvements in GDP, inflation, and unemployment were all positively related to national public support for integration.\textsuperscript{5} In addition, they included variables that captured national and temporal trends: the number of years since 1975 and the timing of entry.\textsuperscript{6} Timing of entry was coded so as to distinguish three groups of members: the original member-states, those who joined in the first expansion in 1975, and those who joined in the 1980s. As in Eichenberg and Dalton (1993), the substantive impact of macroeconomic conditions was small relative to the impact of the temporal and nation-specific variables.\textsuperscript{7}

\textsuperscript{4} For example, a 5-point increase in the percentage of a nation’s exports to EU member-states accounted for almost a 2-point increase in national public support. Similarly, a 5% inflation rate is associated with almost a 10-point decrease in support. In contrast, the United Kingdom dummy variable accounted for more than 40-point difference in support, relative to all other EU members except Ireland and Denmark.

\textsuperscript{5} It is important to note that, unlike Eichenberg and Dalton (1993), Anderson and Kaltenthaler (1996) did not account for any nation-specific fixed effects in their statistical model. To the extent that such fixed effects exist, the results reported in Anderson and Kaltenthaler are biased.

\textsuperscript{6} Vaubel (1994: 171-2) also found a positive relationship between years of membership and public support for integration. Other variables, hypothesized to capture national economic returns from integration, were not consistently related to support for integration.

\textsuperscript{7} In their statistical analysis, timing of entry was associated with a 16 percentage-point difference in national public support for integration. In contrast, GDP growth, inflation and unemployment rates of 5% accounted for differences of only 1-2 percentage points.
In short, previous studies have shown that national variation in public support for integration is due in small part to economic conditions and in large part to nation-specific variables. Unfortunately, there are no theoretical explanations for these nation-specific differences. Dalton and Eichenberg (1993) attributed the impact of national dummy variables to differing “national traditions” of foreign relations and political culture. While such national characteristics may influence support, there is no theory to guide \textit{a priori} predictions about how national traditions will influence variation in public support for integration. Moreover, there is no assurance that the fixed effects captured by the national dummy variables are truly attributable to the national traditions identified by Dalton and Eichenberg rather than other nation-specific factors.

Similarly, Anderson and Kaltenthaler (1996) provided little theoretical explanation for why public support should vary with the timing of entry. Their division of member-states into three groups was based on the assertion that: (a) differences in the level of public enthusiasm for integration and historical circumstances influenced the timing of entry and (b), in turn, these historical differences in levels of national enthusiasm influence more recent levels of national public support for integration (Anderson and Kaltenthaler 1996: 177). They claimed, in other words, that public opinion at the time of entry influences public opinion in the future. This assertion begs the question of why some national publics are historically more enthusiastic about integration than others. As in the case of national traditions, we prefer to account for national differences in support related to timing of entry with a generalizable and testable theory. In other
words, we want to replace the names of countries or groups of countries with the names of theoretically meaningful variables.

In sum, previous research tells us three things about public opinion toward European integration: (1) the public lacks interest in and a sophisticated understanding of integration; (2) national public support for integration varies in systematic ways; and (3) this variation is due in large part to nation-specific factors, particularly those related to the timing of EU entry. Thus, two important puzzles remain. First, how can we reconcile the public apathy toward integration with the observed systematic variation in national public support for integration? Second, how can we account theoretically for national variation in support?

We address these two issues in the following section by developing a theoretical explanation for national variation in public support for integration that assumes citizens lack interest in and sophisticated knowledge of the integration process. Relying on recent research in American public opinion, we first present a model of consistent aggregate-level public opinion that does not require sophisticated individual-level opinion formation. We then present our theoretical explanation for variation in national public support for integration.

A. Low Information Opinion Formation

How can generally uninformed citizens form attitudes about European integration that generate systematic and predictable national differences in public opinion? Scholars have traditionally concluded that such opinion formation is impossible. Early studies of American public opinion claimed that, particularly on foreign policy issues, generally uninformed citizens
lack well-behaved belief systems and render public opinion superficial and transitory (Rosenau 1961; Converse 1964; Neuman 1986). Recent research, however, has challenged this conclusion. Several studies have found that the American public forms consistent collective opinions on foreign policy, despite having limited information and lacking significant interest (Aldrich, et al. 1989; Russett 1990; Page and Shapiro 1992; but see Bartels 1996). Page and Shapiro (1992: xi) summarized their findings as follows:

The main argument is that the collective policy preferences of the American public are predominantly rational, in the sense that they are real—not meaningless, random "nonattitudes"; that they are generally stable, seldom changing by large amounts and rarely fluctuating back and forth; that they form coherent and mutually consistent (not self-contradictory) patterns, involving meaningful distinctions; that these patterns make sense in terms of underlying values and available information; that, when collective policy preferences change, they almost always do so in understandable and, indeed, predictable ways, reacting in consistent fashion to international events and social and economic changes as reported by the mass media; and, finally, that opinion changes generally constitute sensible adjustments to the new conditions and new information that are communicated to the public.

Two aspects of the opinion formation process are central to this argument. First, individual citizens use readily available informational cues and "short-cuts" from their social and political environment to infer their own policy preferences (Brady and Sniderman 1985; Conover and Feldman 1989; Popkin 1991; Kuklinski and Hearly 1994; Huckfeldt, Beck, Dalton and Levine 1995). The mass media, interest groups, and politicians supply the public with cheap information by condensing complicated political information into simple and easily understood messages. Based on these messages, people form beliefs about the implications of particular policies for themselves and their communities. Through this inferential process, many uninformed citizens can
form political opinions that appear informed. Second, any individual-level errors in judgment tend to cancel each other out through aggregation, resulting in coherent collective opinion (Page and Shapiro 1992).

Applying this argument to the EU, we posit that a significant portion of the European public can infer sufficient information from the messages of national political elites and media to generate consistent national opinions about European integration. Thus, we expect aggregate-level support for integration to vary systematically with the tenor of this information about integration. Given that the media and national elites generally focus on the national implications of issues, we expect their messages to reflect national concerns regarding the benefits and costs of integration. Thus, we expect national public support to vary with national differences in the value of European integration.

B. National Interests and Public Support for European Integration

What is the basis for national differences in the value of European integration? We expect the value of integration to vary across nations with the tangible economic and political consequences of integrative policy. The central motivation for unifying Europe was to integrate its members’ economies so as to promote economic exchange and to enhance political stability on the continent. To a great extent, these economic and political goals have been achieved, with varying implications for EU member-states depending on their national political and economic interests. Thus, we expect differences in national benefits from these consequences of integration to account for national-level variation in support for integration.
There are two general areas in which EU membership has important implications for national welfare. First, the EU is a customs union that facilitates trade through the reduction of tariffs and protectionist regulation. Second, the EU promotes political stability—both internal and external—by establishing institutional links between the EU nations, their economies, and their government elites. In the remainder of this section, we specify a set of hypotheses that link EU policy implications to public support.

**EU Policy Implications: Intra-EU Trade**

Given that EU membership constitutes more open trade policy with the other EU nations, we hypothesize that a nation's EU support is related to its benefits and costs from more open trade.\(^8\) The impact of EU trade liberalization on each nation’s economy depends on the structure of its intra-EU trade and the importance of EU trade to its international commerce. Specifically, we hypothesize that a positive relationship exists between a nation's public support for European integration and that nation's trade balance with the other EU nations.\(^9\) A nation's EU trade

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\(^8\) Even though economic theory states that free trade is Pareto optimal, many nations adopt protectionist international trade policies. International Political Economy (IPE) scholars find that nations with superior bargaining power, economic size, and industrial specialization typically obtain a larger portion of the gains from trade. In turn, IPE scholars have identified several practical explanations for protectionism based on political rent-seeking, institutional "stickiness", and national security interests (Katzenstein 1985; Hall 1986; Magee, et al. 1989).

\(^9\) This hypothesis is mercantilist in that imports are detrimental and exports are beneficial to a nation. This mercantilist viewpoint is often adopted as a central premise in Realist analyses of foreign trade policy (Hirschman 1980; Conybeare 1987). Even so, some readers may question why we only investigate this general hypothesis and ignore more-detailed IPE hypotheses derived from the Stolper-Samuelson and Ricardo-Viner models of international trade. Our statistical analysis is restricted to this general hypothesis since data limitations prevent us from differentiating among respondents according to their economic interests in scarce- and abundant-factor industries or import-competing and export-oriented industries.
balance represents a relatively simple objective measure of the (short-run) effects of EU membership on that nation's citizens, industries, and general economic situation. Trade liberalization provides nations with intra-EU trade surpluses with dependable access to markets for their exports while nations with deficits cannot protect their domestic producers from intra-EU competition.

Elites and citizens may be interested in the amount of trade as well as the balance of trade when considering the implications of a more open market among the EU nations. EU membership constitutes an advantage for nations who emphasize intra-EU commerce because mutual liberalization makes trade more dependable and efficient. Therefore, we also hypothesize that a nation's public support for European integration is positively related to the importance of EU trade as part of that nation's overall international trade. This argument is similar to that presented by Katzenstein to explain the persistent support of free trade by small nations. Given the size of their economies, Katzenstein (1985: 43-4) contends that "exporting the costs of [economic] change through protectionist policies is not a viable political strategy" for small states, and that their political autonomy and economic welfare are best served by promoting free trade. As a result, the elites and mass publics in the small European states generally favor free trade over protectionist policies. Unlike Katzenstein, however, our hypothesis focuses on a nation’s international trade orientation rather than its economy’s dependence on EU trade. We believe that public support in a particular nation increases as a greater portion of its international trade is conducted with other EU nations regardless of the overall importance of international trade to its
In other words, we expect the citizens in nations whose international trade is more sharply oriented towards intra-EU commerce to more strongly support integration as a means of promoting free trade among EU member-states.\(^\text{11}\)

**EU Policy Implications: External Stability**

The purpose for European integration is not purely economic. European cooperation is also motivated by a Western European desire to institutionalize economic links between its states in order to neutralize the divisive nationalism that led to two world wars. Jean Monnet described this motivation:

> The purpose of European integration is to eradicate the spirit of superiority, of domination which has driven the various nations of Europe, brought wars, and almost the end of Europe and can again, if not destroyed, bring the end of the world...It is the beginning of organizing for peace. Otherwise, domination will be pursued, and with nuclear arms, that is the destruction of humanity (May 9, 1950 press conference, Paris).

The creation of the European Coal and Steel Community (ECSC), the initial step toward European integration, demonstrated the security motivation for European economic cooperation. One of the sources of controversy after World War II was the future of German coal and steel

\(^\text{10}\) An alternative, though similar, hypothesis is that citizens’ support for European integration increases with their national economy’s reliance on EU trade, in particular, and international trade, in general. This alternative hypothesis is actually more consistent with Katzenstein’s argument. We thank Jeff Frieden for suggesting this theoretical distinction. We prefer our trade-orientation hypothesis, however, since it accounts better for temporal variation in national public support due to shifts in international trade toward greater commerce among the EU nations. Even though the statistical analysis presented in the paper only considers our trade-orientation hypothesis, we investigated the validity of the trade-reliance hypothesis in auxiliary analyses discussed briefly below.

\(^\text{11}\) Note that Eichenberg and Dalton (1993) made a similar prediction about public support for integration, but they based their trade hypothesis on neo-functionalist theories that contend greater international interactions promote more positive public attitudes toward international cooperation.
resources. France was in control of the coal-rich Saar basin, to which Germany had territorial claims. In addition, the industrial Ruhr area of Germany was under international control. For security reasons, the French—along with other Europeans—wanted to keep this militarily valuable coal and steel out of German hands. For economic reasons, West Germany needed these resources to rebuild its economy. To resolve this dilemma, Robert Schuman proposed a supranational authority to oversee the tariff-free movement of coal, steel, and iron within the territories of France, West Germany, and other European states. The proposal was realized with the creation of the ECSC. Subsequent economic integration—the creation of the EEC and EU—have further reduced the ability of EU member-states to use exclusionary or restrictive economic policies for military advantage.

National publics differ in the benefits that they derive from this enhanced external security, depending on how sensitive they are to its primary consequence: the reduced chance of the recurrence of traditional military conflict on the continent. For nations that suffered great physical and human losses (e.g., West Germany and France) from the previous military conflicts on the continent (i.e., the two world wars), this reduced threat war represents a substantial increase in their security. For nations that suffered smaller losses (e.g., Spain and Ireland), the value of this reduced threat of conflict is of relatively less importance to their national security. Consequently, we expect EU citizens to differ in their security-based evaluations of integration according to the losses their nation endured from previous military conflict among the EU members. Specifically,
we posit that a national public's support for integration is positively related to its death toll in World War II.\textsuperscript{12}

In addition, the salience of external security concerns in citizens' appraisals of integration should decline over time. While the “German question” may have weighed heavily on the minds of Europeans immediately following the war, the threat of German military hostility towards the EU members decreased substantially over the subsequent decades. Indeed, one of the clearly positive results of European integration has been the absence of military hostility between the EU member-states over the past fifty years. As a result, “armed conflict between any two [members] does not now appear to be remotely possible” (Nugent 1994: 7). Thus, while security concerns may vary across national publics depending on their national experience with the world wars, for all national publics the relevance of these security concerns should decay over time. In order to investigate this hypothesis, we include an interaction variable in our statistical model that captures any change in the influence of external stability concerns over time.

Note that this hypothesis differs fundamentally from the security hypothesis proposed by Eichenberg and Dalton (1993). Eichenberg and Dalton posited that EU citizens’ security concerns regarding East-West conflict influence their support for integration. We do not share this contention because European integration does not necessarily influence the likelihood of East-West conflict. The North Atlantic Treaty Organization (NATO) and the United Nations were designed explicitly to deal with such conflicts. Consequently, we do not expect citizens' appraisals

\begin{footnote}
\textsuperscript{12} We use the total number of World War II casualties—military and civilian—per capita, as a proxy for the costs of war. Tilly (1992: 201) lends credence to this measure by stating that, "With the 20th century, battle deaths underestimated more and more the damage done by war. The bombing and shelling of civilian settlements destroyed increasing numbers of non-combatants, not to mention their means of livelihood."
\end{footnote}
of integration to reflect their concerns over East-West conflict or their general national security. Eichenberg and Dalton (1993) presented evidence in this regard, showing that the level of East-West conflict is not statistically related to support for European integration.

**EU Policy Implications: Internal Security**

European integration has also influenced the internal security of member-states by promoting the consolidation of democratic capitalism. Prior to World War II, most nations in Western Europe had developed democratic governments and capitalist economic systems. Many of these regimes suffered from both economic and political instability, particularly during the 1930s. This instability, in conjunction with the eventual fall of these regimes to the Nazis, fostered political opposition to democratic capitalism after the war (Lindberg and Scheingold 1970: 6). On one side, Communist parties espoused the replacement of capitalist economies with command economies, by whatever political means necessary. On the other side, extreme right wing parties emerged that rejected pluralistic democracy in favor of authoritarian direction of the government and economy. Consequently, the post-war democratic-capitalist regimes of Western Europe faced political challenges from both ideological extremes of the political spectrum (Urwin 1972: 154).

European integration has reinforced democratic capitalism, as Urwin (1972: 279) argues:

After 1945 the extent of wartime damage and dislocation, the dearth of economic resources, and the advent of the Cold War had posed to parliamentary democracy in Western Europe a challenge as grave in its implications as the Nazi threat. To a large extent this threat had been countered, especially in France and Italy with their large Communist parties, by the massive influx of American financial aid and by the developments of the plans for European and Atlantic communities.
The creation of an internal market in the EU was designed to promote economic predictability and growth for its members, based on a capitalist system. This economic certainty and growth enhanced EU nations' abilities to respond to the economic demands of their electorates. For one, the post-war economic growth financed the dramatic expansion of the welfare state in Western Europe. These welfare policies addressed many of the economic concerns of the working class, a traditional base of support for Communist parties (Hoffman 1983: 35; Slater 1983). In addition, economic integration reduced the chance of economic downturn and rising unemployment, which are related to increased support for anti-democratic right-wing parties. In short, the contribution of European integration to domestic economic conditions helps its member-states reduce domestic political opposition to democratic capitalism.

The extent to which national publics value this benefit depends on the current threat to political stability. Nations with strong political opposition to democratic capitalism should value integration more than nations with only weak political opposition. We theorize that the existence of an internal political threat to democratic capitalism mobilizes support for integration among those citizens who favor democratic capitalism. The stronger is the threat, the greater public support mobilizes in response to it. Thus, we expect public support for integration to increase with the level of domestic opposition to democratic capitalism. Implicit in this hypothesis is the assumption that the majority of citizens support democratic capitalism.  

13 Of course, in the democratic regimes of the EU member-states, if the opposition ever reached a majority, the status quo political-economy would likely change and our hypothesis would not apply.
As an indicator of opposition, we use electoral support for parties that oppose democratic capitalism. This includes all communist parties (opponents of capitalism) and a select group of parties that oppose democratic governance. The selection of these parties is discussed in Appendix B. To capture temporal changes in the intensity of public opposition to democratic capitalism in each nation, we use a three-election moving average of support for these opposition parties.

Note that if our theory is correct, external and internal stability concerns should have influenced national publics’ enthusiasm for, and hence their governments decisions about, participating in European integration at its initial stages. Consequently, nations’ decisions to join the European Coal and Steel Community—the original component of the EU created in 1951—should reflect the relative strength of their concerns about external and internal stability.\(^\text{14}\) We consider this hypothesis by comparing the security concerns of the nations that originally joined the ECSC with those that did not.

For each potential ECSC member, Figure 1 plots the average vote percentage for political parties opposing democratic-capitalism from 1945-1951\(^\text{15}\) and the percent of population killed in World War II. A clear distinction exists between the nations that initially did and did not join. The original ECSC members—France, West Germany, Italy, Belgium, the Netherlands, and Luxembourg--had at least 0.56% of their population killed in World War II and an average opposition vote share of at least 7.6% during 1945-1951. In sum, Figure 1 suggests that

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\(^{14}\) Note that we do not use the decision to join the EEC since the EEC was created by the ECSC members, which only considered the United Kingdom as a possible new member. Thus, regardless of other nations’ interests in joining the EEC, the opportunity was not available to them.

\(^{15}\) The parties included were all communist parties, the Reich Party (Germany), and the MSI (Italy).
differences in external and internal stability concerns are what distinguish the nations that originally joined the ECSC from those that did not. Moreover, these results provide a possible explanation for Anderson and Kaltenthaler’s (1996) finding that order of entry is related to support for integration. Since we expect these security concerns to shape both nations’ decisions to enter the ECSC and their publics’ enthusiasm for integration, they provide the theoretical connection between national entry into the ECSC and public support for integration in contemporary periods. We will examine more rigorously the relationship between EU public support and internal and external security concerns in the ensuing analysis.

[Figure 1 about here]

II. Data and Methodology

In order to investigate the hypothesized relationships discussed above, we have pooled time-series data for 1975-1992 across EU member-states. The data set is constrained on both ends by availability. Thus, we only analyze public opinion in the first twelve member-states, although we would expect the hypotheses to apply to the three more recent members as well. In constructing the data set, we utilized the Eurobarometer surveys to measure a nation's public support for European integration and OECD data to measure a nation's international trade interests in EU membership. The data set is discussed in more detail in Appendix A.

The dependent variable in our statistical analysis, EU Support, is the mean level of support for integration among a nation's citizens in a particular year. Consistent with Gabel (1998b) and
Gabel and Palmer (1995), this variable is constructed from responses to two questions in the Eurobarometer surveys. These questions are worded as follows:

Generally speaking, do you think that (your country's) membership in the European Community (Common Market) is a good thing, neither good nor bad, or a bad thing?

In general, are you (very much/to some extent) for or against efforts being made to unify Western Europe?

Responses to these questions are coded so that higher values correspond with greater support. To calculate the dependent variable, each individual's coded responses are summed, standardized by the maximum possible total, and then multiplied by one hundred, so that the individual-level measure of EU support ranges from zero to one hundred. A nation's mean level of public support is then obtained by taking the average standardized sum for all of a nation's respondents in a particular year. As a result, a nation's mean level of public support can range from zero—minimum support among all its respondents—to one hundred—maximum support among all its respondents. Gabel (1998b) showed that this variable is a better proxy for support for actual integrative measures than any other common survey measure. A full discussion of the coding of the dependent variable is provided in Appendix A.

The explanatory variables in our analysis are designed to measure the effects of international trade interests, and external and internal stability concerns. The measurement of the explanatory variables is described in Appendix A. Descriptive statistics are given in Table 1. The reader should note that the dependent variable has a mean of 73.1 and a standard deviation of 10.4. This implies that most of the variation in EU Support is within 11 points of the mean. Since
each explanatory variable accounts for only a fraction of this variance, the magnitude of the estimated coefficients may appear small relative to the 100-point scale.

[Table 1 about here]

The paper's statistical models are estimated with annual time-series cross-section data. An advantage of the TSCS design is that it allows us to examine the cross-temporal and cross-national variations in EU Support simultaneously. A disadvantage of the TSCS design is that the error term typically violates the Gauss-Markov assumptions of constant variance and no autocorrelation. In the presence of heteroscedasticity and autocorrelation, parameter estimates generated by OLS are consistent but not efficient. More importantly, OLS estimates the wrong variance-covariance matrix for the parameter estimates, and the OLS estimate of the error variance may be biased. Hence, OLS standard errors produce misleading inferences.

In the present context, we suspect that both groupwise heteroscedasticity and first-order autocorrelation plague our data, so OLS is problematic. One approach to these problems is to use feasible generalized least squares (FGLS). FGLS estimators have the attractive statistical properties of consistency and asymptotic efficiency. The asymptotic efficiency of FGLS estimators, however, may not carry over to small samples because of variability introduced by the estimated error covariance matrix (Greene 1997: 512-13). Moreover, Monte Carlo experiments reveal that the FGLS procedure tends to underestimate the coefficient standard errors in small samples (Freedman and Peters 1984; Beck, et al. 1993).

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16 Stimson (1985) offers an introductory discussion of the methodological issues associated with TSCS models; Greene (1997: 648-69) provides more sophisticated treatments of the econometric techniques used to estimate these models.
An alternative approach is to use OLS to estimate the model’s parameters, since these estimates are still consistent, and then evaluate these parameter estimates with robust standard errors that have been corrected for possible misspecification in the error covariance matrix (see Greene 1997: 652-3). Several authors have suggested specific applications of this approach that assume different forms of misspecification (White 1980; Arellano 1987; Newey and West 1987; Beck and Katz 1995). The general logic of the “robust standard error” approach is to derive a consistent estimator of the appropriate asymptotic covariance matrix for the OLS coefficients, \( \text{Var}[b] = (X'X)^{-1}X'VXX'(X'X)^{-1} \).

This requires a consistent estimator of \( X'VX \), as White (1980) shows. Applications differ in how they specify the error covariance matrix, \( V = E[ee'] \). For TSCS data, the most general case is,

\[
V = \begin{bmatrix}
\sigma_{11}\Omega_{11} & \sigma_{12}\Omega_{12} & \ldots & \sigma_{1n}\Omega_{1n} \\
\sigma_{21}\Omega_{21} & \sigma_{22}\Omega_{22} & \ldots & \sigma_{2n}\Omega_{2n} \\
\vdots & \vdots & \ddots & \vdots \\
\sigma_{n1}\Omega_{n1} & \sigma_{n2}\Omega_{n2} & \ldots & \sigma_{nn}\Omega_{nn}
\end{bmatrix},
\]

where \( \Omega_{ij} \) are \( T \times T \) submatrices, \( n \) is the number of units, and \( T \) is the number of time periods (see Greene 1997: 651-2). In this most general case,

\[
X'VX = \sum_{i=1}^{n} \sum_{j=1}^{n} \sigma_{ij} X'_{i} \Omega_{ij} X_{j}.
\]

By imposing the assumption of no serial correlation among disturbances from different time series (i.e., \( E[e_{it}e_{js}] = 0 \) for all \( i \neq j \) and \( t \neq s \)), expression 3a reduces to,

\[
X'VX = \sum_{i=1}^{n} \sigma_{ii} X'_{i} \Omega_{ii} X_{i} + \sum_{i=1}^{n} \sum_{j \neq i} \sigma_{ij} X'_{i} X_{j}.
\]
By also imposing the assumption of no serial correlation in the error term within cross-sections (i.e., $E[\varepsilon_i \varepsilon_{it}] = 0$ for all $i$ and $t \neq s$), expression 3b reduces to,

$$\sum_{i=1}^{n} \sum_{j=1}^{n} \sigma_{ij} X_i' X_j = \sum_{i=1}^{n} \sum_{j=1}^{n} \sigma_{ij} X_i' X_j. \quad (3c)$$

This specification allows for groupwise heteroscedasticity and cross-sectional correlation, and is assumed in the panel-robust estimator suggested by Beck and Katz (1995),

$$\text{Var}[b] = \left( \sum_{i=1}^{n} X_i' X_i \right)^{-1} \left( \sum_{i=1}^{n} \sum_{j=1}^{n} \sigma_{ij} X_i' X_j \right) \left( \sum_{i=1}^{n} X_i' X_i \right)^{-1}. \quad (4)$$

Before applying this estimator, any serial correlation in the data must be removed. Beck and Katz (1995) suggest including a lagged dependent variable on the right-hand side of the regression model. The advantage of this approach is that it explicitly models the time-series process. This approach also assumes, however, that the dependent variable is autoregressive, implying that it is determined by lagged exogenous variables as well as lagged disturbances. If the

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17 The TSCS design structures our expectations about what form of heteroscedasticity can arise. Therefore, we implicitly assume that the covariances among contemporaneous disturbances do not vary over time, i.e., $E[\varepsilon_i \varepsilon_{jt}] = \sigma_{ij}$ for all $t$, $i$, and $j$. In turn, the specification of groupwise heteroscedasticity incorporated into expression 3c is more restrictive than a pooled version of the estimator proposed by White (1980). When analyzing TSCS data, however, the Beck-Katz estimator is superior to the White estimator (e.g., t-statistics constructed with it have greater power), since it imposes assumptions consistent with the pooled design. In stronger words, the White estimator is inappropriate for TSCS data since it ignores the pooled nature of the data.

18 Consistent estimates of $\sigma_{ij}$ are derived with $\hat{\sigma}_{ij} = \frac{e_i' e_j}{T}$, where $e_i$ is a vector of OLS residuals for unit $i$. Expression 3c indicates why Beck-Katz standard errors are generally larger than their OLS counterparts. In accounting for groupwise heteroscedasticity, the weighting of elements in $X'X$ suggests the possibility that panel-robust standard errors could actually be smaller. The second term in the middle part of expression 3c, however, will always be positive definite, in practice. Hence, unless little cross-sectional correlation exists, the main-diagonal elements of $X'VX$ will be larger than the corresponding elements of $\sigma^2 X'X$. 24
researcher has theoretical reasons to doubt the autoregressive nature of the dependent variable, correcting for autocorrelation may be a more appropriate approach. We apply both approaches in the statistical analysis of EU Support presented below.

Accounting for serial correlation, shifts the focus of the regression analysis from levels to quasi-differences. Regardless of whether the researcher includes a lagged dependent variable or corrects for autocorrelation, the transformed regression model explains variation in the dependent variable’s rate of change rather than its level. But what if the researcher wants to explain cross-sectional differences in the level of the dependent variable? Is it possible to use OLS regression coefficients estimated with data plagued by autocorrelation as well as groupwise heteroscedasticity and cross-sectional correlation?

In response to these questions, we propose an alternative to the technique suggested by Beck and Katz (1995). As stated earlier, in the presence of autocorrelation, heteroscedasticity and cross-sectional correlation, OLS parameter estimates are consistent but the corresponding covariance matrix is wrong. Following the general logic of the robust standard error approach, we need to derive a consistent estimator of the appropriate asymptotic covariance matrix. Expression 3b above provides the specification of $X'VX$ that allows for the possibility of autocorrelation, groupwise heteroscedasticity and cross-sectional correlation. Newey and West (1987) have devised a covariance estimator that is consistent in the presence of autocorrelation and heteroscedasticity. While the Newey-West estimator is designed for the case of a single time series, adapting it to the TSCS context is relatively straightforward. The first right-hand-side term in expression 3b captures any autocorrelation within each time series. By applying the
Newey-West method to each time series separately, we can obtain a consistent estimate of this term. The second right-hand-side term in expression 3b can be estimated consistently in the same manner described above for the Beck-Katz technique. In sum, adapting the Newey-West method to the TSCS context produces the following estimator,

\[
\text{Est. Var}[b] = \left( \sum_{i=1}^{n} x_i' x_i \right)^{-1} \left( \hat{Q} + \sum_{i=1}^{n} \sum_{j \neq i} \hat{\sigma}_{ij} x_i' x_j \right) \left( \sum_{i=1}^{n} x_i' x_i \right)^{-1},
\]

where \( \hat{Q} = \sum_{i=1}^{n} \hat{\sigma}_{ii} x_i' x_i + \sum_{i=1}^{n} \sum_{l=1}^{L} (1 - \frac{1}{L+1}) e_{i-t} e_{i-t-l} (x_{i-t} x_{i-t-l} + x_{i-t-l} x'_{i-t}). \)

As with the Beck-Katz estimator, \( \hat{\sigma}_{ii} \) and \( \hat{\sigma}_{ij} \) are derived with OLS residuals (see note 19).

Notice the similarity between this alternative robust estimator and the Beck-Katz estimator. The only difference is the incorporation of the second term in \( \hat{Q} \), which accounts for the possibility of autocorrelation. As with the Newey-West estimator, the researcher must choose \( L \) in advance so that autocorrelations at lags longer than \( L \) are trivial in magnitude. Unfortunately, there is little theoretical guidance in making this determination (see Davidson and MacKinnon 1993: 613). In the analysis presented below, we choose \( L = 6 \).

One criticism of using OLS coefficients evaluated with robust standard errors is that in the presence of heteroscedasticity, OLS may produce less accurate parameter estimates than FGLS.

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19 We have slightly modified the Newey-West estimator that is applied to each time series. Consistent with the TSCS design, we have imposed a groupwise-heteroscedasticity specification on the Newey-West estimator. In a pooled version of the original Newey-West estimator,

\[
\hat{Q}_p = \sum_{i=1}^{n} \sum_{t=1}^{T} e_{i-t} x_{i-t} x'_{i-t} + \sum_{i=1}^{n} \sum_{l=1}^{L} (1 - \frac{1}{L+1}) e_{i-t} e_{i-t-l} (x_{i-t} x'_{i-t-l} + x_{i-t-l} x'_{i-t}).
\]
On the basis of this criticism, Greene (1997: 667-8) suggests another approach in which the parameter estimates are derived with partial-FGLS that accounts for groupwise heteroscedasticity and evaluated with a corrected coefficient covariance matrix that is robust to the possibility of cross-sectional correlation. As with the approach suggested by Beck and Katz (1995), any serial correlation in the data must be removed by correcting for autocorrelation or including a lagged dependent variable in the regression model.

In Greene’s approach, the parameters are estimated with the traditional FGLS estimator (i.e., weighted least squares with a lagged dependent variable or after serial correlation has been removed from the data) and evaluated with the following covariance matrix,

\[
\text{Est. Var}[\hat{b}_{\text{FGLS}}] = \left( \sum_{i=1}^{n} \frac{1}{\hat{\sigma}_{ii}} \mathbf{X}_i \mathbf{X}_i \right)^{-1} \left[ \sum_{i=1}^{n} \sum_{j=1}^{n} \frac{\hat{\sigma}_{ij}}{\hat{\sigma}_{ii} \hat{\sigma}_{jj}} \mathbf{X}_i \mathbf{X}_j \right] \left( \sum_{i=1}^{n} \frac{1}{\hat{\sigma}_{ii}} \mathbf{X}_i \mathbf{X}_i \right)^{-1},
\]

where \( \hat{\sigma}_{ii} \) and \( \hat{\sigma}_{ij} \) are derived with OLS residuals (see note 19). One-step FGLS generates consistent parameter estimates. ML estimates can be obtained by applying iterated FGLS and scanning the admissible range of autocorrelation parameters (i.e., \(-1 < \rho < 1\)). Alternatively, ML estimates can be derived directly by maximizing the following log-likelihood function,

\[
\ln L_{\theta} = -0.5 \left[ \ln(2\pi) + \ln \sigma_i^2 - dfo_i \left( \ln(1 - \rho^2) - (1 - \rho^2) \frac{u_i^2}{\sigma_i^2} \right) + (1 - dfo_i) \left( \frac{u_i - \rho u_{i-1}}{\sigma_i^2} \right)^2 \right],
\]

where the first right-hand-side term is a pooled version of the White estimator that allows for a more general form of heteroscedasticity. In contrast to expression 8, the estimator defined in expression 5 assumes that \( E[e_i^2] = \sigma_i^2 \) for all \( t \).
where \( u_t = y_t - \beta x_t \) and \( dfo_t \) is a dummy variable denoting the first observation \( (t = 1) \) for cross-section \( i \). This log-likelihood function incorporates the Prais-Winsten transformation for the first observation and generalizes the error variance in the log-likelihood function for first-order autocorrelation by allowing for groupwise heteroscedasticity.

In this section, we have described four alternative methods of panel-robust estimation: (1) OLS coefficients derived with the inclusion of a lagged dependent variable and evaluated with Beck-Katz standard errors; (2) OLS coefficients evaluated with panel-adapted Newey-West standard errors; (3) ML coefficients derived with an AR(1) error process and homoscedasticity imposed and evaluated with Beck-Katz standard errors; and (4) ML coefficients derived with an AR(1) error process and allowing for groupwise heteroscedasticity and evaluated with Greene standard errors for weighted least squares. The choice of which method to use is not necessarily trivial since the four methods impose different assumptions about the structure of the causal process and hence may produce coefficients with different levels of precision. Unfortunately, the information needed to choose the most appropriate and precise method is unobservable. Thus, researchers must choose among alternative methods based on their theoretical expectations about the nature of the dependent variable and error term.

In discussing each of the methods above, we highlighted the structural assumptions imposed to derive them and thereby suggested some applied circumstances that might motivate their use. If the researcher has theoretical reasons to believe that the causal process is

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\(^{20}\) We obtained the ML estimates reported in the paper by applying the direct approach but we also tried the iterated FGLS method and obtained essentially the same coefficients (i.e., coefficient differences were so small that they could be attributed to rounding between iterations).
autoregressive, at least in the error term, methods 1 and 3 are more appropriate. It is rare in the social sciences, however, for theory to indicate the presence of only an autoregressive error process. Generally, the autoregressive nature of a dependent variable applies to all causal factors, not simply the unobserved error term. Thus, the researcher will rarely face a situation in which method 3 (i.e., correcting only for autocorrelation) is clearly more appropriate than method 1. As Beck and Katz (1995) suggest, if the researcher suspects an autoregressive causal process, the advantage of explicitly modeling the time-series process with a lagged dependent variable generally outweighs the possibility of misrepresenting the process as something more than serial correlation in the error term.

For some social scientists, the use of time-series data alone, even in a pooled design, is sufficient justification for explicitly modeling the causal process as (entirely) autoregressive and, in the present context, adopting method 1. In some applied settings, however, cross-sectional variation represents a larger portion of total variation and is as, if not more, important as explaining temporal variation. Under these circumstances, method 1 has the disadvantage of shifting the focus of the analysis from the dependent variable’s level to its rate of change, since much of the cross-sectional variation is captured by the lagged dependent variable. This disadvantage can be partially addressed by interpreting the autoregressive model as an infinite distributed (or geometric) lag model and calculating the long-run effects (or multipliers) of the independent variables (see Greene 1997: 796-9). Method 1 still obscures, however, the significance of variables that primarily explain cross-sectional variation in the level of the
dependent variable. Thus, in applied settings in which theory and scholarly interest focuses on cross-sectional variation, method 2 is more appropriate.

Finally, if theory indicates the existence of groupwise heteroscedasticity, it seems appropriate to apply method 4 at least as a check on the robustness of the results produced by another method (most likely 1 or 2). When the TSCS data is composed of national-level variables, for example, there are generally convincing theoretical reasons, such as differences in historical traditions and electoral rules in the field of comparative politics, to believe that the error variance varies across nations. Hence, in such contexts, we recommend strongly the use of method 4.

In the specific context of analyzing EU Support, we are interested in accounting for cross-national differences in the level of support and we are not convinced that mean national support is autoregressive (i.e., a function of lagged values of the explanatory variables). Moreover, we believe that external security concerns, measured by War deaths per capita and War deaths*\log(time), primarily account for cross-national differences in the level of support for integration. In an autoregressive specification of EU Support, the explanatory variables largely account for differences in the rate of change in support so we expect external security concerns to have little relevance in such a specification. Thus, a specification that explains variation in the level of support is more appropriate for investigating the relevance of external security concerns. In sum, our theoretical expectations suggest that the most appropriate approach to analyzing EU Support is method 2: OLS coefficients evaluated with panel-robust Newey-West standard errors (even though in the presence of severe groupwise heteroscedasticity, method 4 might produce
more precise parameter estimates). Despite these theoretical considerations, we apply all four panel-robust methods in our statistical analysis in order to consider how the empirical evidence changes across estimation methods.

III. Statistical Results

Tables 2 and 3 present the results of our statistical analysis. The results in column 1 of Table 2 and columns 1 and 3 of Table 3 are OLS coefficients derived with the inclusion of a lagged dependent variable and evaluated with Beck-Katz standard errors. The results in column 2 of Table 2 and columns 2 and 4 of Table 3 are OLS coefficients evaluated with panel-adapted Newey-West standard errors. The results in columns 3 and 4 of Table 2 are ML coefficients—the former derived with homoscedasticity imposed and evaluated with Beck-Katz standard errors, and the latter derived with groupwise heteroscedasticity and evaluated with Greene standard errors.\(^{21}\)

[Tables 2 and 3 about here]

A limited number of national dummy variables are included in each of the models to avoid biased estimates due to fixed effects.\(^{22}\) We first estimated each model without national dummies

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21 For brevity, we do not report the nation-specific error variances for the ML model in column 4 of Table 2. These error variances range from 2.28 for the Netherlands to 20.15 for Great Britain. Other than those for Greece, Spain and Portugal, the error variances are all significantly different from zero at the 5% level. Using a LR test derived with the log-likelihood values for the models in columns 3 and 4 of Table 2, we can reject the null hypothesis of homoscedasticity at better than the 1% level. The LM statistic of groupwise heteroscedasticity is 45.1, constructed with the OLS residuals from the second model in Table 2 (Greene 1997: 655), which gives the same test result.

22 In order to assess robustness, we re-estimated the models in columns 1 and 2 in Table 2 eleven times, “dummying out” the observations for a different EU member-state with each trial. This procedure enabled us to evaluate how sensitive the coefficients are to the “exclusion” of a particular nation’s observations. We found that the statistically significant coefficients in Table 2 generally changed by only 10-30%, and never by more than 48%, due to the “exclusion” of a nation’s observations. Moreover, none of these
and then, based on the national residual means, included national dummies so that none of the national residual means in the final model were significantly different from zero at the 10% level. Model specifications that include and exclude national dummy variables largely support the same inferences.

It is important to emphasize that the results for the model specifications including and excluding national dummy variables are very similar in two significant aspects. First, the addition of national dummy variables changes the measures of overall model fit very little. When the national dummy variables are excluded, the $R^2$ values drop to .906, .795, .807, .907 and .807 for the OLS models in columns 1 and 2 of Table 2 and 2-4 of Table 3. The LR statistics drop to 396.3 and 427.4 for the ML models in columns 3 and 4 of Table 2. Thus, there appears to be little systematic variation in national support for integration that is not captured by our explanatory variables. Second, the coefficients for the explanatory variables in the models that exclude national dummy variables are either unchanged or larger in magnitude and statistical significance than in the models that included national dummy variables. The only exception are the coefficients for EU trade balance per capita, which are smaller in magnitude and statistical significance in the models that exclude national dummy variables. The regression results for specifications without national dummies and the national residual means for all models are available from the authors as part of a supplemental appendix.

Consistent with Beck and Katz (1995), t-statistics constructed with the panel-robust standard errors in Tables 2 and 3 are smaller than their OLS and one-step FGLS counterparts.
For the five core explanatory variables (i.e., excluding the constant, national dummy variables and lagged dependent variable), we calculated the average attenuation in t-statistics. Compared to one-step FGLS, the core variables’ panel-robust t-statistics are 31.3, 6.3, 18.0 and 15.8 percent smaller, respectively, for the four models in Table 2. Compared to OLS, these panel-robust t-statistics are 10.4, 23.3, 36.8 and 32.2 percent smaller. In sum, statistical inferences based on one-step FGLS and OLS standard errors are optimistic.

Assessing the hypotheses:

Starting with the security hypotheses, the regression models provide strong evidence consistent with our theory. War deaths per capita—the proxy for external security concerns—is positively related to public support for integration. Also, the effect of this security concern declines over time. The interaction term between war deaths per capita and time is negative, showing a decay in the positive influence of external security concerns on support for integration. Based on the results in column 2 of Table 2, for instance, external stability concerns in West Germany, the Netherlands, and France increased public support for integration by 20.3, 8.5, and 5.5 points respectively in 1975, but by only 10.4, 4.4 and 2.8 points respectively in 1992. As stated above, we did not expect the first model to produce results consistent with the external security hypothesis since the lagged dependent variable accounts for cross-national differences in the level of EU Support.

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the Irish observations were “excluded” from each model.
As for internal security, the results indicate that as political opposition to democratic capitalism grows, support for integration increases. In the second model, for instance, a one standard deviation (one-SD) increase in electoral support for parties opposed to democratic capitalism is associated with a 6.4 point increase in national public support for integration. The marginal effect of *Anti-democratic capitalism vote* does not vary much with the method of panel-robust estimation, given that the smaller coefficient in the first model only reflects the impact of internal security concerns on temporal variation in *EU Support*.

The results are also consistent with our hypotheses concerning intra-EU trade. National economic benefits from intra-EU trade are strongly related to public support for integration. Intra-EU trade balance is positively correlated with support. Using the coefficient in the second model, a one-SD increase in *EU trade balance per capita* ($44.83 per capita) results in a 2.2 point increase in *EU Support*. While the other panel-robust estimation methods generate smaller coefficients for *EU trade balance per capita*, they are still significantly significant at the 5% level. We also discover that a nation's public support is strongly correlated with its emphasis on intra-EU trade. According to the second model, a one-SD increase in *EU trade concentration* (10.5) increases *EU Support* by 4.6 points. As with *Anti-democratic capitalism vote*, the different panel-robust methods estimate similar marginal effects of *EU trade concentration*.

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23 As stated above, an alternative hypothesis here is that *EU Support* increases with the national economy’s reliance on intra-EU trade. We tested this alternative hypothesis with two sets of auxiliary analyses (presented in a supplemental appendix). In the first set, we replaced *EU trade concentration* with *intra-EU trade/GDP*. In these models, *intra-EU trade/GDP* had a statistically significant positive effect, even though it displayed less explanatory power than *EU trade concentration*. In the second set, we added *intra-EU trade/GDP* to the regression models in Table 2. In these models, *intra-EU trade/GDP* did not achieve statistical significance at the 10% level with all four t-statistics being less than 1 in magnitude.
Controlling for rival hypotheses:

To examine the robustness of these relationships, we re-estimated the first and second models in Table 2 including control variables for alternative explanations for national public support for integration.\textsuperscript{24} The first two columns of Table 3 present models including a time trend variable designed to control for the effect of political socialization on public support for integration. Several scholars assert that public support has increased over time as a result of political socialization (Inglehart, et al. 1991; Anderson and Kaltenthaler 1993). Similar to our external and internal stability hypotheses, the political socialization hypothesis specifies why public support is greater in the original members than in the later members. By controlling for length of membership, these models ensure that the coefficients on the stability variables are not capturing the effect of political socialization over time.

As shown in Table 3, the coefficients for \textit{Length of membership} are statistically significant at the 10\% level but have different signs. Contrary to theoretical expectations, the estimated marginal effect of length of membership is negative in the OLS model that includes \textit{Lagged EU Support}. This finding suggests that support for integration declines over time since \textit{Lagged EU Support} captures cross-national differences in the initial level of support (e.g., due to the

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\textsuperscript{24} In the same manner, we re-estimated the third and fourth models in Table 2. For brevity, we do not report these auxiliary results here but have made them available as part of a supplemental appendix. The general pattern is the same as that for the comparison of OLS models in Tables 2 and 3. In the auxiliary ML models, \textit{Length of membership} and the macroeconomic variables prove insignificant, and their inclusion reduces slightly the statistical significance of the other variables but does not alter substantially the inferences drawn from the models.
original/later member distinction). Consistent with this inference, the coefficient on *Length of membership* is positive in the second model where it accounts for cross-national as well as cross-temporal differences in the level of support. This finding seems counter-intuitive, especially since political socialization over time is generally assumed to have a positive effect on public support for integration. Meanwhile, all of the other coefficients in the first and second models of Table 3 remain consistent with those reported in Table 2.

In columns 3 and 4 of Table 3, we include control variables for national macroeconomic conditions. Inglehart and Rabier (1978) and Eichenberg and Dalton (1993) posited the variation in national public support for integration is related to objective national economic conditions. Specifically, Eichenberg and Dalton hypothesized rising support with improving national economic conditions: real gross domestic product, unemployment, and inflation. To control for this potential influence on support, we include control variables for these macroeconomic indicators.

The results presented in the third and fourth models of Table 3 indicate that these control variables have little effect on our hypothesized relationships. The parameter estimates for our explanatory variables remain largely unchanged. The coefficients for the macroeconomic indicators, however, are generally inconsistent with the expectation that support increases with improving national economic conditions. In the fourth model, inflation attains a statistically significant negative effect, consistent with expectations, but this result does not hold for the alternative panel-robust estimation methods.²⁵

²⁵ Eichenberg and Dalton (1993) also obtained a significant negative coefficient for inflation.
Discussion:

So far we have assessed the strength of the estimated relationships in Table 2, but have not fully considered how cross-national and cross-temporal differences in economic and political concerns have affected support for integration. Notably, the results of our statistical analysis demonstrate that the original/later member distinction can be attributed to differences in the importance of external and internal stability concerns at the inception of the EU. Furthermore, the empirical evidence also indicates that there has been a decline over time in the contribution of external and internal stability concerns to public support in the original members. Not only has the salience of external stability concerns decreased, the electoral strength of the parties opposed to democratic capitalism have converged across the EU member-states. Except for Italy, the relatively high levels of opposition in the original members has declined recently. And, the electoral strength of the Communist Party has increased in several of the later members—Greece, Spain, Portugal, and Denmark. In addition, economic interdependence in the EU has increased over time, thereby reducing national differences in the importance of intra-EU trade. Consequently, the original/later member distinction has softened over time as public support in the original members has became more dependent on economic factors and less dependent on political factors.

Overall, the statistical models perform extremely well. Despite using a limited number of national dummy variables in our analysis, the models in Table 2 explain about 90% of the
variation in public support for European integration. More importantly, over 75% of the variation in EU Support is attributable to the variables measuring national economic and political concerns rather than to the national dummy variables. The models represent a significant improvement over past empirical research in that they precisely explain cross-national differences in public support for integration. Moreover, they account for the original/later member distinction in support with specific comparative relationships rather than with “catch-all” national effects.

IV. Concluding Remarks

The paper proposes and applies a new panel-robust method for analyzing time-series cross-section data: OLS coefficients evaluated with panel-adapted Newey-West standard errors. The principal advantage of this method is that it does not require the researcher to specify a time-series process. In the specific context of explaining national public support for European integration, this new method largely produces the same statistical inferences as the methods proposed by Beck and Katz (1995) and Greene (1997). One important exception to this generalization, however, is the estimated effects of external stability concerns. The inclusion of a lagged dependent variable in the OLS model, as suggested by Beck and Katz, accounts for much of the national differences in the level of public support for integration. Hence, external stability

\footnote{Eichenberg and Dalton (1993) report an $R^2$ of 0.69 for their FGLS model of EU support, which included dummy variables for six of the eight nations in their data set. $R^2$ values from FGLS and OLS models, however, are not precise counterparts so comparing them is not entirely appropriate.}

\footnote{Sometimes applied circumstances provide little guidance in choosing a time-series process. Beck and Katz (1995) suggest, for example, that the researcher impose a common, rather than nation-specific, autoregressive or AR(1) error process when the number of observations is small (as is often true for TSCS}
concerns appear to have an inconsequential effect on EU Support. Beck and Katz’s method does not account, however, for national differences in the “initial” level of EU Support (i.e., the level at \( t = 0 \) or, in the specific context of our data, \( t = 1975 \)). Given that such differences are of substantive interest to students of EU politics, their method provides an incomplete explanation.

From a substantive perspective, the paper's statistical analysis has general implications for the study of public opinion and specific implications for understanding the process of European integration. First, our analysis complements recent research on consistent public opinion in the American context (Page and Shapiro 1992; Russett 1990; Aldrich, et al. 1989). By demonstrating that national-level public support for European integration varies in meaningful ways across the twelve EU members, our analysis provides evidence of the generality of consistent public opinion in a comparative context. Moreover, given the general lack of public interest and information concerning EU issues, it is intriguing to find that public support for integration can be predicted so accurately with objective measures of national economic and political benefits from EU membership.

Second, the present study contributes to our understanding of European integration by explaining cross-national and cross-temporal variation in public support for integration. Public opinion, particularly through national referenda, represents a crucial constraint on future integration. Consequently, in order to explain and predict the dynamics of integration, we must understand why national publics differ in their support for integration. The results of our statistical analysis identify several economic and political factors that explain a large portion of the data in comparative politics). But in the present case, the null hypothesis of a common AR(1) error process
variation in support for European integration among the publics of the first twelve member-states. We find that the national publics increase in support for integration with their degree of economic reliance on intra-EU trade and political reliance on EU membership for maintaining internal and external stability.

These empirical findings have important implications for the future of European integration. First, our results indicate that one of the primary sources of higher public support--external stability concerns--is declining in salience over time. Consequently, in the future, the distinction between original and later members may become less clear, and support for integration in the original members may become more sensitive to economic interests and thus less stable. Second, our results suggest that the EU's increasing economic interdependence should lead to increased public support for integration. It is important to note, however, that current German economic interests in Eastern Europe and the former Soviet Union could reduce this interdependence. Given Germany's large share of total intra-EU trade, a diversion of German trade to non-EU nations could reduce the extent of economic interdependence in the EU and thereby dampen public support for integration.

Finally, the analysis indicated that two common explanations for variation in national public support for integration are dubious. When the effects of our explanations are accounted for, neither national economic conditions nor length of membership influence support for integration as expected. Indeed, length of membership is negatively related to support, which is counter to previous findings. A possible reason for this result is that length of membership serves can be rejected at the 1% level with a LR test ($\chi^2[10] = 31.9$).
simply to distinguish the relatively supportive original members from the rest. We offer several theoretical reasons why the original members exhibit greater support and thus capture directly what length of membership might measure if examined alone.
Appendix A: Coding of Dependent and Explanatory Variables

Our statistical analysis utilizes pooled time-series data for the EU nations from 1975-1992. The dependent variable, *EU Support*, was constructed with Eurobarometer surveys and the intra-EU trade variables with OECD data. The Eurobarometer is conducted twice a year. Yet, our data set uses an annual time series rather than a semi-annual one, since the OECD trade data is annual. Whenever possible, we use respondents from both Eurobarometers in a particular year to construct the relevant variables. The EU nations during this period are Belgium, Denmark, France, Ireland, Italy, Luxembourg, the Netherlands, the United Kingdom, and West Germany with the additions of Greece in 1981 and Portugal and Spain in 1986. We treat Belgium and Luxembourg as a single nation, however, since the OECD combines their trade data.

*EU Support* is the mean level of support among a nation's respondents in a particular year. For Belgium and Luxembourg, respondents from both nations are used in the calculation of *EU Support* but those respondents are weighted differently since small nations are over-sampled in the Eurobarometer surveys. For further details on the construction of *EU Support*, see Gabel and Palmer (1995).

The regression models of *EU Support* presented in Tables 2 and 3 include a variety of explanatory variables. The measurement of these variables is described below:

*EU trade balance per capita* is a nation's intra-EU trade balance divided by its population, measured in U.S. dollars per capita. Source: OECD, *Foreign Trade by Commodities*.

*EU trade concentration* is the percent of a nation's total international trade (exports and imports) that is conducted with the other EU nations. Source: OECD, *Foreign Trade by Commodities*.

*WWII deaths per capita* is the percent of a nation's 1939 population that was killed in World War II. Sources: Sivard 1985; Snyder 1982; and Keegan 1989.

*WWII deaths* *log(time)* is an interaction variable equal to the product of *WWII deaths per capita* and the natural log of the number of years since 1974.

*Anti-democratic capitalism vote* is a three-election moving average of a nation's vote percentages.
for parties that oppose democratic capitalism, as defined in Appendix B. Sources: Mackie and Rose 1995; *Keesing's Archives*; and the *Europe Yearbook*.

*Length of membership* is the number of years that a nation has been an EU member.

*Real GDP* is gross domestic product adjusted by purchasing power parities and measured in 1990 US dollars. Source: OECD, *Main Economic Indicators*.

*Inflation* is the annual percentage change in consumer prices. Source: OECD, *Main Economic Indicators*.

*Unemployment* is the rate of unemployment, standardized when available. Source: OECD, *Main Economic Indicators*.

**Appendix B: Coding of Parties Opposed to Democratic Capitalism**

In order to construct a proxy for internal security concerns, we identified two types of parties opposed to democratic capitalism. First, we included all communist parties and related parties in Ireland (Workers’ Party), Greece (Coalition of the Left), Portugal (United Democratic Left), Italy (Proletarian Union, Continuous Struggle, and Workers’ Vanguard), and Spain (United Left). Second, we included extreme right-wing parties (ERPs) that were anti-democratic capitalism. Based on Ignazi and Ysmal (1992) and Ignazi (1992), we added the following ERPs: National Front (France), NPD and REP (West Germany), EPEN (Greece), MSI (Italy), PDC (Portugal), AP, FNs, and Falange (Spain).

Note that we did not include ERPs that attract votes primarily for reasons other than opposition to democratic capitalism. For example, neither the Vlaamsbloc in Belgium nor the Center Democrats in the Netherlands espouse a clear opposition to democratic capitalism (Mudde 1994). Note also that, although regional or separatist parties may threaten the stability of governing institutions, we exclude them from our measure. The reason is that European

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28 Note that parties were not selected on the basis of their positions toward European integration. Some of the ‘opposition’ parties we identify have been supporters of integration, e.g., the Italian Communist Party (PCI). We simply chose parties that represent political threats to democracy and/or capitalism, at least as perceived by the public. Thus, while the PCI has been supportive of integration and has participated in democratic institutions, mainstream public apprehension about the party’s commitment to capitalism and democracy continued throughout the post-war period (see Spotts and Weiser 1986: chapter 3).
integration does not help the regime withstand such political opposition. Recall that our hypothesis is that European integration, by reducing opposition to democratic capitalism, enhances the stability of the governing institutions. Since these parties are not contesting democratic capitalism, we do not expect the threat of opposition from these parties to influence support for integration.
References


Organization for Economic Cooperation and Development (various years). *Main Economic Indicators.* Paris: OECD.


<table>
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<td></td>
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Notes: Panel-robust standard errors are reported in parentheses: Beck-Katz in the first and third columns, panel-adapted Newey-West (with L=6) in the second column, and Greene for weighted least squares in the fourth column. The maximum likelihood coefficients were estimated using a common AR(1) error process and the Prais-Winsten transformation of the first observation for each nation. *p<0.05; **p<0.01
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<td>( R^2 )</td>
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Notes: Regression coefficients were estimated by least squares. Panel-robust standard errors are reported in parentheses: Beck-Katz in the first and third columns, and panel-adapted Newey-West (with \( L=6 \)) in the second and fourth columns. *\( p<0.05 \); **\( p<0.01 \)
Divided Opinion, Common Currency:  
The Political-Economy of Public Support for EMU

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Much of the political economy literature on European integration has focused on how the distributional consequences of economic integration structure political conflict over integration. Specifically, studies have traced how economically advantaged and disadvantaged segments of society have responded politically so as to further their interests in economic integration. Underlying these studies is an assumption that at least some citizens in these different segments of society identify their economic interests with regard to integration and then adopt political attitudes and behavior consistent with these interests. Thus, this assumption is central to a many political economy explanations for European integration.

Recent research on public support for European integration supports this assumption. While EU citizens are generally disinterested in and uninformed about European integration, there is evidence that variation in their support for integration is consistent with their economic interests from market liberalization. The focus of this presentation will be on whether this holds true for public support for Economic and Monetary Union (EMU). That is, does public support for EMU vary with the distributional impact of the creation of common currency?

The attached paper provides some background on public opinion and support for European integration. The paper focuses on variation in support across national publics and over time. The presentation will focus on individual-level variation in support for EMU. Related empirical results and discussion will be forthcoming.