Short-Term Fluctuations in U.S. Voting Behavior, 1896–1964*

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I. Introduction

This study is an attempt to employ some simple statistical models, motivated by certain assumptions about voting akin to those discussed by Downs and others, in an attempt to explain short-term fluctuations in the division of the national vote for the U. S. House of Representatives, over the period 1896–1964. The models will yield quantitative estimates of the impact of economic conditions on congressional elections, and of the effects of incumbency and presidential "coattails" as well.

The notion that a vote represents a decision or rational choice between alternatives is an important theme in democratic theory. However, this rationality hypothesis has proved to be difficult to test empirically, particularly with survey data, from which most of our recent knowledge of individual voting behavior is drawn. The present study is an attempt to put a modified form of the rationality hypothesis to a different and in some respects more direct test than is readily possible with survey data.

The analysis bears directly on the substantive question of the relationships between economic conditions and U. S. national election results. National economic decisions frequently involve such considerations as how much price stability must be sacrificed in order to achieve a specified growth rate, and so forth, so clearly, quantitative knowledge of the electoral consequences of varying mixes of growth, price stability, and unemployment is relevant for an incumbent administration wishing to maximize its own chances for reelection; conceivably, such knowledge may also be relevant to the task of achieving a socially optimal mix of policies. Moreover, investigation of these various relationships may also provide a basis for developing new methods of long-range election forecasting to supplant current techniques, based on polls. The specific relationships used in the present study are too aggregative to be of great forecasting value in themselves, but our results do suggest that a more disaggregated and detailed model of the same general type may be of interest in this respect.

Finally, the extent to which congressional election results depend on economic and other external conditions, which cannot be effectively controlled by any particular congressman or campaign organization, is a question of general interest to students of campaigning and elections, and might also have implications for our understanding of the operation of the division of powers in the Federal government.

Several previous studies have considered the question of the effect of economic conditions upon election outcomes, and the principal results should be briefly described here.

W. A. Kerr examined the correlations of various economic indices with the "conservative" presidential vote (defined as the Whig vote prior to 1856, and the combined Republican-Prohibition vote in subsequent years) over a series of elections, in an attempt to test the hypothesis that prosperity increases the conservative vote. The rank-order correlations with a variable identified only as "Index of per capita realized national income (adjusted by cost of living)", averaged over the election year and three preceding years, was .19, over the period 1897–1940. The correlation with "cost of living (corrected for century trend)," similarly averaged, was .17 (for 1837–1936), and with a "wholesale price index," .29 (for 1861–1940). More explicit definitions of the variables or data sources were not given, and most of the other economic series used are not measures of "prosperity" in any usual sense of the term. All in all Kerr's results lend only modest support, at best, to his basic hypothesis.

Pearson and Meyers considered a somewhat different hypothesis, that the "public tends to vote for the continuation of administrations that have been in power during prosperous times and to vote against the incumbent administration when depression marks the approach of election time." They used a "general price level" as their measure of overall prosperity, and found that the presidential candidate of the incumbent party was defeated in 11 of the 13 presidential elections from

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2 Though see V. O. Key, Jr., The Responsible Electorate (Cambridge: Belknap Harvard, 1966).


1828 to 1924 which occurred in years of "declining or low prices," and was victorious in 16 of the 18 elections held when prices were "rising or high." The authors were not explicit, however, about what data were used, or about exactly how election years were classified into their two categories.

L. H. Bean considered essentially the same hypothesis, using an index of general business activity "based on Cleveland Trust Co. and Am. Tel. and Tel. indexes." He compared changes in the index, between October of the election year and October two years earlier, with changes in the House membership of the incumbent administration. During nineteen congressional elections between 1854 and 1954 the index declined over the two-year period, and in fifteen of these cases the incumbent party suffered a net loss in House membership. Such losses occurred in all nine midterm elections, but the pattern held for only six of the ten presidential elections considered. Bean concluded that the president's personality and other factors partially offset the impact of economic conditions in presidential years, and that the effect of economic decline was most pronounced in midterm elections. The possibility of the opposite effect—that economic upturn would increase the incumbent's House seats—was also considered, but Bean concluded that this type of effect did not occur with any regularity. The data on which this conclusion rested, however, were presented in a rather fragmentary fashion.

Bean's findings were sharply disputed by Wilkinson and Hart, who also used an index of general business activity (at one point identified as the Cleveland Trust Company index, and at others as "Ayer's index.") Changes in the August–September–October average of the index between the election year and the previous year were found to be completely uncorrelated with the percent of popular presidential vote for the party in power (for presidential elections from 1844 to 1948), and with changes in the House membership of the party in power (for congressional elections from 1856 to 1948); in each case the product-moment correlation was reported as .000. The authors concluded that the "alleged relationship [between prosperity and political victory] is not supported by the facts."72

The studies cited above were longitudinal, in the sense of analyzing aggregate national election outcomes over a long series of elections. A second group of studies employed cross-sectional data, by counties or states, over a shorter series of elections. The study by Tibbits was intended to test the thesis that economic decline or upturn would hurt or help the incumbent party in the Congressional elections of 1882 and 1884. The "incumbent" vote was defined separately in each congressional district, according to which party held the district prior to the election. This "incumbent" vote was then aggregated over all the districts in each state, and various procedures, whose rationale is not very clear, were then applied to this statewide "incumbent" vote to "correct" for trend and the influence of presidential elections. Although Tibbits regarded the thesis as verified, the data he presents are not very convincing, and in any event the definitions and procedures employed make it difficult to compare his results with those of other studies.8

Gosnell and Coleman investigated the relation between changes in the Democratic presidential vote and changes in economic conditions in some sixty-five Pennsylvania counties, for the elections of 1928, 1932, and 1936. Their economic variable was an index "derived from the amounts of wages and salaries paid in manufacturing enterprises and the value of principal crops ... [weighted by] the relative importance of agriculture and industry ... [as measured by] the number of persons gainfully employed in each group [for each county]."73 Changes in this index between 1928 and 1932 were found to correlate at −.231 with changes in the Democratic presidential vote over the same period, and the equivalent correlation from 1932 to 1936 was +.236. (Some of the partial correlations, holding constant the rural-nonrural or ethnic-nonethic composition of the counties, were somewhat higher in magnitude, on the order of .3 to .5). These results support the basic thesis that economic fluctuations affect the incumbent president's vote, though the correlations are small and account for only around 6% of the variance of the dependent variable.

A similar study, using counties from several states, was done by Ogburn and Coombs. They used several economic indices, the most relevant of which was based on "average wages of those engaged in manufacturing and in retail and wholesale trade ... [and] the average value of farm per person engaged in farming ... [weighted by] the proportion of the population living in towns and

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74 Ibid., p. 334.
on farms (in each county).[9,10] Changes in this index between 1932 and 1936 were correlated across counties with corresponding changes in the Roosevelt vote, with mixed results. The reported correlations were: Illinois, −.15; Indiana, −.23; Ohio, −.03; Pennsylvania, −.04; Kansas, +.39; Iowa, +.01; Nebraska, +.32; and California, +.24.11 Sizeable positive relationships, consistent with the "incumbency" thesis, were found in only three states, while substantial negative relationships were found in two others. These findings, and indeed most of the other results obtained in the study, do not correspond well with those of Gosnell and Coleman, and create the overall impression of no systematic relationship between changes in prosperity and the Roosevelt vote.

In the most recent study by Rees, et al., the relationship between the Republican congressional vote, and "state insured unemployment, as a percent of covered employment," and "net income per farm" were investigated in each of forty-one states during the seven congressional elections from 1946 to 1958. Each of these three variables was expressed in terms of deviations from its thirteen-year (or seven-election) state averages and simple bivariate tabulations were used to see whether below-average unemployment or above-average farm income was associated with an above-average Republican vote, and conversely. The authors found some modest association with unemployment (189 or 289 cases from 1946–58, and 162 of 235 from 1948–58, supported the thesis), but none with farm income. The unemployment relationship was strongest in the twenty-three northeastern and midwestern states.12

Finally, there have been several studies based on poll data, rather than election returns. Clark examined the relationship between monthly series on Roosevelt's popularity as President, as measured by Gallup poll results, and national income. He reports a correlation of .57 between these two variables (both expressed as deviations from a linear trend), over the 57-month period from August 1935 to April 1940. (When the popularity figures are computed separately for various economic classes, the correlations range from .78 to .95 for various income classes, and is .33 for persons on relief, over the shorter February 1937–April 1940 period.)13 Duranti cites several journals.


II. The Model

To some extent, at least, an individual's vote in a national election represents a choice or judgment between alternative governing "teams." The precise interpretation of this choice depends on the decision rule one assumes voters to follow. One might picture the voter as a rational, information-processing individual who proceeds by collecting information of various kinds—party platforms and policy pronouncements, legislative voting records, and perhaps expert or authoritat-

tive opinions on these matters. Such a voter analyzes this information in the light of his own self-interest and decides which party presents the "best" package of positions. He then votes accordingly. This view of the voting decision appears in classical democratic theory, and has been subjected to empirical tests in various voting studies (where it generally does not fare very well).

In a more realistic setting, however, a voter—even a rational, self-interested voter—may not find it practical or efficient to proceed in this manner. For example, there may be no relevant party or "team" platforms to compare; platforms may (indeed, usually do) concentrate on desired ends rather than specific policy proposals; voters may not feel qualified to make a confident a priori assessment of the relative merits of positions on subtle technical issues, or they may recognize that platforms are in no sense binding commitments. Other information, such as the detailed legislative voting records of individual candidates, may be very costly to acquire and analyze, and of only limited relevance to the issues at stake in a national campaign.

These considerations suggest that a more relevant decision rule for voters would be based on readily available information. The past performance of the incumbent party in particular gives some indication of what it would do if returned to office, and of the effectiveness of its policies and personnel.

With respect to short-term variations in voting behavior, we shall, therefore, assume that a decision rule of the following type is operative: if the performance of the incumbent party is "satisfactory" according to some simple standard, the voter votes to retain the incumbent governing party in office to enable it to continue its present policies; while if the incumbent's performance is not "satisfactory," the voter votes against the incumbent, to give the opposition party a chance to govern.

On the basis of these assumptions, a simple statistical model can be developed as follows: consider a sequence of elections, numbered 1, 2 \ldots T, which are contested by two political parties, A and B. For election t, let

\[ y_t \] be party A's share of the two-party vote

\[ \delta_t \] be +1 if A is incumbent

\[ -1 \] if B is incumbent

\[ \Delta_t \] be the difference between the actual performance of the incumbent party during its just-ending term in office, as measured by some index such as per capita income, and the performance "expected" of it by the electorate.

\[ u_t \] be a random disturbance, distributed in some partially specified fashion \[ \alpha, \beta, V \] be parameters

With these definitions, a suitable model is of the form

\[ y_t = V + \delta_t[\alpha + \beta \Delta_t] + u_t \quad (2.1) \]

Here, the random variable \( u_t \) is interpreted as the net effect of factors not considered explicitly in the model, such as the personalities and campaign tactics of individual candidates running on the party ticket, or the state of foreign affairs. The quantity \( V \) is the "normal" long-run average vote for party A; this is the vote A would receive in the absence of incumbency or random effects, and in this sense is a measure of the underlying basic partisanship of the electorate. The parameter \( \beta \) is a measure of the effect of the discrepancy \( \Delta_t \) between actual and expected performance of the incumbent; from our basic hypothesis, we should expect \( \beta \) to be positive. Finally, \( \alpha \) is a measure of the institutional advantage (or disadvantage) of being incumbent.

For an election in which A is incumbent, A will receive its base vote \( V \), plus the net incumbency advantage \( \alpha \), plus the quantity \( \beta \Delta_t \), which reflects the effect of its performance in office; when B is incumbent, then the vote for A will be \( V - \alpha - \beta \Delta_t \), since the latter two terms associated with the incumbent now add to B's vote. The other factors embodied in the random term \( u_t \) also affect the vote, so that the observed relationships will actually be scatters of points around these lines.

If we could observe \( \Delta_t \) over a series of elections, and if we could make suitable assumptions about the effects of the other factors embodied in the latent variable \( u_t \), it would be possible to fit the model to the data, and estimate its parameters. In particular, if \( u_t \) satisfied the standard Gauss-Markov assumptions, we could estimate \( \alpha \) and \( \beta \) from an ordinary least-squares regression. The intercept term would be an estimate of \( V \), and the coefficients of \( \delta_t \) and \( \delta_t \Delta_t \) would be estimates, respectively, of the advantage of incumbency (\( \alpha \)) and of the impact of the incumbent's economic performance (\( \beta \)). It is not possible, however, to serve voters' expectations directly, however, so we must supplement the model with additional hypotheses. A reasonable and convenient hypothesis is that expectations about year \( t \) are formed on the basis of experience during the preceding year, \( t-1 \). With respect to income, for example, we might assume that income is "expected" by the average member of the electorate to grow at some constant rate \( r \) from year to year, and that the percentage discrepancy between actual and expected income \( Y_t \) and \( \hat{Y}_t \) in year \( t \) serves as \( \Delta_t \) in (2.1). Thus, using the relation \( \hat{Y}_t = (1+r)Y_{t-1} \) to express this year's expected income in terms of last year's actual income, the relation becomes
\[ y_t = V + \delta_t[\alpha + \beta(Y_t - \bar{Y}_t) + \bar{Y}_t] + \epsilon_t \]
\[ = V + \delta_t \left[ \alpha + \frac{\beta}{(1+r)} \left( \frac{Y_t - Y_{t-1}}{Y_{t-1}} - r \right) \right] + \epsilon_t. \]

This relation, however, contains an extra parameter, \( r \), which can be neither observed nor estimated directly. More precisely, the three parameters \( \alpha, \beta \) and \( r \) are not identified; unless we have other prior information or other types of data, it is not possible even in principle to disentangle their separate effects. If we simply regress \( y_t \) on \( \delta_t \) and \( \delta_t(Y_t - Y_{t-1})/Y_{t-1} \), the coefficients of these variables will be estimates of \( \alpha - r\beta/(1+r) \) and \( \beta/(1+r) - r\beta/(1+r) \), respectively, rather than the original parameters \( \alpha \) and \( \beta \). For plausible values of \( r \)—say if income is expected to grow at not more than 5% per year—the bias in the case of \( \beta \) is not too serious, since the \( -r\beta/(1+r) \) term will result in a negative bias of less than 5% of the true value of \( \beta \). In the case of \( \alpha \), however, the bias may be more serious. For \( r = .05 \) and \( \beta = .5 \), the bias \( -r\beta/(1+r) \) will be on the order of \( -0.02 \). If we think of \( \alpha \) as the institutional advantage accruing to the incumbent (arising from patronage, better access to the media, and so on), it seems unlikely that this could be more than a few percent of the total vote; hence a bias of two percent is substantial, relative to the true value of \( \alpha \). Thus, the estimated coefficient of \( \delta_t \) in the regression above must be interpreted as the net advantage of being incumbent, which increases with institutional advantages, but decreases when voters have high expectations with respect to income increases.

In applying this framework to the U.S., further complications arise from the fact the model, as described so far, applies to a simple Downsian type of quasi-parliamentary electoral system, in which a voter must decide only between two candidates, who are running primarily as anonymous members of opposing party “teams” (rather than on the basis of their own personal records or policy views) and who are contesting a single office. In the U.S., however, voters must decide separately between candidates for several offices (president, senator, congressman). Because of the lack of party discipline and the fact that different offices are contested at different times and in different constituencies, candidates who are nominally members of the same party frequently do not campaign at all like members of the same electoral “team.”


The present analysis will focus on the vote for the House of Representatives, since of the races for national office, House contests come closest to the Downsian case of relatively anonymous candidates competing as members of a common party team. Clearly, individual candidates often do not behave or campaign as loyal members of a party team, and indeed sometimes campaign on the basis of their past or intended future independence of their party’s leadership. However there is considerable evidence, at least for recent times, to suggest that such factors count for very little in votes. For example, Stokes and Miller conclude that in the 1958 mid-term election, straight party loyalty was by far the major factor in the vote for congressmen; they also found that well over half the voters did not know either House candidate in their district, and that of those who knew something about either candidate, only a negligible percent mentioned legislative or policy matters. Although individual races may deviate from the overall pattern, in general it seems that most Congressional candidates appear to most voters simply as Democrats or Republicans, and not as clearly defined personalities with their own policy views and records; and hence, variations in the overall popularity of the parties should be a major factor in producing short-term fluctuations in the Congressional vote.

The definition of the “incumbent” party also requires some care, since because of the division of powers it is possible for one party to control either or both houses of Congress, while the other controls the presidency; this situation in fact occurred in eight of the thirty-five national elections from 1896 to 1964. In three of these elections—1896, 1920, and 1932—the division of control was a transitory phenomenon which arose only because the president’s term happened not to coincide with the beginning of the electoral tide against his party; the same party had formerly controlled all three institutions of government, and then progressively lost control of all three in the mid-term and following presidential elections. In these cases it seems quite appropriate to regard the party occupying the presidency as the incumbent. The election of 1912, in which control was also divided, was not used in the analysis for other reasons, described below. The remaining four cases—the elections of 1948, 1956, 1958, and 1960—are more complicated, and no definition of “the” incumbent party seems fully satisfactory. Even when control is divided, however, the president has greater control over the machinery of government, and normally has greater initiative in

policy matters than does the majority party of either house of congress. The president also is
the most prominent member of his party, and the
most prominent elected official, while the identity
of the party controlling the legislature is much
less well known. (According to Stokes and Miller,
even in the 1958 election, when the Democrats
had controlled the House for four years, less than
half the electorate was aware of that fact.) For
these reasons, we will take the party which con-
trols the presidency as the incumbent party.
While most of the votes cast in national elec-
tions go to the two major parties, various minor
parties also compete and regularly manage to win
a small fraction of the vote. During the period
from 1896 to 1920, minor parties usually received
from four to six percent of the total congressional
vote, and substantially more in some elections
(notably in 1912 and 1914, with the Republican-
Progressive split); from 1920 to 1942 the minor-
party vote was typically from two to four percent,
and since then it has generally run less than two
percent of the total. Variations within the above
approximate limits have often been substantial
relative to the inter-election variation in the
major-party vote. Hence we should try to incorpo-
rate the minor-party vote into the analysis, rather
than lose information by simply ignoring it. Since
we have been interpreting short-term fluctuations
in the division of votes as expressions of satisfac-
tion or dissatisfaction with the performance of the
incumbent party, it seems appropriate to interpret
the minor-party vote as part of the anti-incum-
ental party, and to count it along with the major
opposition party. Hence (except where otherwise
stated), when we speak of the "Republican" share
of the vote, this quantity will consist of the Republic-
ian share of the total vote when the Republicans
are the incumbent party, and of the Republican
plus minor-party share when the Democrats are
incumbent. The "Democratic" vote will be treated
similarly. This procedure was followed except for
the election of 1912, when the incumbent Republic-
ian party itself was split. There seems to be no
satisfactory way of interpreting the sizeable Prog-
ressive vote in that year, so for this reason the
1912 election has been omitted from the analysis.
A final complication which must be con-
trolled is the possibility of "coattail" effects from high-
level races. If a candidate in a more prominent
and higher-level race runs well ahead of his ticket
because of his personal attractiveness and cam-
paigning ability, some of the extra votes he wins
may carry over to lower-level congressional races,
simply because of the intellectual and physical dif-
culty of voting a split ticket. Our model should
take this type of coattail effect into account, since
we are primarily interested in that portion of the
vote which reflects the strength of the party
"team" as a whole, and not in votes which reflect
the personal attractiveness of individual can-
didates. Because of their variety, it would be diffi-
cult to consider separately the possible effects of
individual statewide races on the congressional
vote. Evidence from other sources indicates that
statewide factors in general (including coattail
effects from statewide races) have little effect on
congressional races, compared to national or local
factors, and even if such effects were
important in individual states, the overall effect on
the total congressional vote would be averaged
over several dozen statewide races (where a strong
Democratic gubernatorial candidate in one state
might be offset by an attractive Republican can-
didate in another), so that the final net effect
would be much smaller. In any event, we shall
ignore statewide races, and consider only the pos-
sible influence of presidential coattails.
In a mid-term election, where there is no presi-
dential contest, our model is of the form
\[
y_t^r = \beta_1 X_{1t} + \beta_2 X_{kt} + \cdots (2.2a)
+ \beta_k X_{kt} + u_t
\]
where \(y^r\) is the Republican share of the congres-
sional vote (the minor-party vote being treated as
described above), and where for example \(X_{1t} = 1\),
to give the intercept term \((V'\) in equation (2.1)),
\(X_{kt}\) is +1 if a Republican is president and -1 if
a Democrat is, \(X_{kt}\) is \(X_{kt}\) times the percent
increase in per capita income, and so forth. For elections
in presidential years, we have two relations,
\[
y_t^p = \beta_1 X_{1t} + \cdots + \beta_k X_{kt}
+ u_t + v_t
\]
\[
y_t^c = \beta_1 X_{1t} + \cdots + \beta_k X_{kt}
+ u_t + \gamma v_t, (2.2b, c)
\]
where \(y^p\) and \(y^c\) are the Republican shares of the
vote for president and congress, respectively. The
same exogenous variables \(X_{kt}\) appear in both rela-
tions, with the same coefficients \(\beta_i\). The random
disturbance \(u_t\), which appears in both equations
also, represents the net effect of unmeasured fac-
tors, such as foreign events, which affect the pop-
ularity of the party "team" as a whole. The presi-
dential equation (2.2b) also contains a second
disturbance \(v_t\), which represents the effect of the
specific candidates and campaign tactics in the
presidential race, and in the congressional equa-

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tion (2.2c), it is assumed that a certain portion $\gamma$ (where $0 \leq \gamma \leq 1$) of this presidential effect is also carried over into the congressional vote. Thus $\gamma r_t$ is the size of the coattail effect in election $t$.\(^2\)

If coattail effects were negligible, and if the disturbance $u_t$ satisfied the Gauss-Markoff assumptions, it would be possible to obtain efficient, unbiased estimates of the $\beta_i$ in (2.2) by simply regressing the Congressional vote $y^C$ on the exogenous variables $X_{it}, \ldots, X_{it}$. However if coattail effects are sizeable, then because of the resulting heteroskedasticity the regression estimates would not be efficient, nor would their estimated standard errors or the associated significance tests be accurate; furthermore there would be no way of estimating the coattails parameter $\gamma$, which is of some interest in itself. To circumvent these difficulties the disturbances have been assumed to be normally (and independently) distributed, and we use maximum likelihood methods for estimating the models; the procedure is described in detail in Appendix A. Briefly, the likelihood function is maximized in a stepwise fashion, by assuming a value for the coattails parameter $\gamma$, then obtaining conditional estimates of the other parameters and a conditional maximum of the likelihood function $L(\gamma)$, and then repeating with alternative values of $\gamma$ until the likelihood is near its maximum. By trying enough values of $\gamma$ (and spacing them closely enough) the maximizing value $\hat{\gamma}$ can be approximated as closely as desired. In obtaining the results reported below, a relatively coarse grid with a spacing of .05 was used, so that the trial $\gamma$-values were $.05, .1, .15, \ldots, .95$. (Except that when $\hat{\gamma} < .05$, a finer grid with a .01 spacing was used.) For any assumed value of $\gamma$ the conditional estimates of the remaining parameters are obtained by what is in effect a least-squares regression, where the dependent variable is the congressional vote $y^C$ in mid-term observations, and is a weighted sum of $y^C$ and $y^P$ in presidential-year observations, the weights being functions of the coattails parameter $\gamma$. With respect to hypothesis testing, for certain purposes—in particular for testing $\gamma$ or for testing the composite hypothesis that certain subsets of variables are not significant—likelihood ratio tests were performed. In assessing the precision of individual $\beta_i$'s, I have simply reported the standard errors as calculated according to the usual least-squares regression formula, for the regression corresponding to the maximizing value of $\gamma$. (A better procedure would have been to obtain asymptotic standard errors from the second derivatives of the likelihood function, but this was not done for computational reasons.) The $R^2$ and Durbin-Watson statistic for the regression corresponding to $\hat{\gamma}$ are also reported. The former is in effect the ratio of "explained" variance to the total variance less the variance attributable to the presidential disturbance $\nu$, and as such, gives a general idea of the overall goodness of fit. The Durbin-Watson test for serial correlation of the disturbances is not strictly appropriate here, both because the residuals are not calculated from a simple least-squares regression, and also because the observations are not entirely consecutive, since certain elections were omitted; however, the statistic is still informative concerning the possibility of serial correlation, and is reported in the tabulated results.

### III. Results

The variables used for estimation were as follows:

- $\delta$, the incumbent index, $= +1$ if the incumbent President is Republican, $-1$ otherwise.
- $Y_t$, $Y_{t-1}$, monetary income, $= \text{per capita personal income (or prior to 1919, adjusted GNP)}$ in current dollars, for the year of the election ($Y_t$) and for the preceding year ($Y_{t-1}$).
- $P_t$, $P_{t-1}$, prices, $= \text{consumer cost-of-living index, 1929 = 100, for the current and preceding year.}$
- $R_t$, $R_{t-1}$, real income, $= \text{monetary income deflated by the cost-of-living index, or } Y_t/P_t \text{ and } Y_{t-1}/P_{t-1}$.
- $U_t$, $U_{t-1}$, unemployment, as a fraction of the civilian labor force, for the current and preceding year.
- $T$, $= \text{time, where } T = 0 \text{ in 1896, } T = 1 \text{ in 1898, } \ldots, T = 34 \text{ in 1964.}$

More precise definitions of these variables, and sources, can be found in Appendix B.

As mentioned earlier, the election of 1912 was not used because of the difficulty in interpreting the large progressive vote in that year. The wartime elections of 1918, 1942 and 1944 have also been dropped, since wartime conditions and controls would substantially distort the meanings of our price and income series. The remaining thirty-one elections from 1896 to 1964 were used in obtaining the results reported below.

The simplest version of the model (2.2) would be to use fluctuations in real income to explain...
fluctuations in the division of the vote; hence, taking the congressional equation (2.2c) as representative of the system (2.2a,b,c) we would have

\[ \gamma^c = \beta_0 + \delta \left[ \beta_1 + \beta_2 \left( \frac{R_{t} - R_{t-1}}{R_{t-1}} \right) \right] + u + \gamma v \]

A more elaborate version of the model is obtained by adding a term involving \( T \) to allow for a trend in partisanship, and other terms to measure the independent effects of price and unemployment fluctuations, e.g.

\[ \gamma^c = \beta_0 + \beta_1 T + \delta \left[ B_2 + \beta_2 \left( \frac{R_{t} - R_{t-1}}{R_{t-1}} \right) + \beta_3 \left( \frac{P_{t} - P_{t-1}}{P_{t}} \right) + \beta_4 (U_{t} - U_{t-1}) \right] + u + \gamma v \]  

(3.5)

Intermediate versions of these models (equations (3.2) and (3.4)) are implicitly defined below. Prices and incomes have grown severelfold over the period, so changes in these variables are expressed in relative terms; unemployment, in contrast, has shown no particular trend, so its changes are expressed in absolute terms. (In fact, alternative expressions were also tried for the unemployment variable, including relative changes and absolute levels; all yielded essentially equivalent results.) Finally, we could replace the expressions involving real income \( R \) by similar expressions involving monetary income \( Y \), to investigate the effects of price and monetary-income changes separately; this is done in the equations (3.3) and (3.6), implicitly defined below.

Our basic hypotheses lead us to expect, a priori, that increases in income help the incumbent party, so the coefficients of terms involving \( R \) and \( Y \) should be positive. Rises in prices and unemployment should hurt the incumbent, so that coefficients of terms involving \( P \) and \( U \) should be negative. It is not clear what the sign of the coefficient of the incumbency variable should be, since it is affected in opposite directions by institutional advantages and income expectations (and may also be affected by expectations concerning prices and unemployment).

The results obtained are presented in tabular form in Table 1. Each column of the table corresponds to an alternative form of the model (such as equations (3.1) and (3.5) above), while in each row are the coefficients of the terms at the left end of the row (provided the equation contains such

### Table 1. Models and Data

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<th>Equation</th>
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<td>(3.1)</td>
</tr>
<tr>
<td>1</td>
<td>.510</td>
</tr>
<tr>
<td>( \delta )</td>
<td>-0.0005</td>
</tr>
<tr>
<td>( \beta_2 )</td>
<td>.503</td>
</tr>
<tr>
<td>( \beta_3 )</td>
<td>(.100)</td>
</tr>
<tr>
<td>( \beta_4 )</td>
<td>-.161</td>
</tr>
<tr>
<td>( \beta_5 )</td>
<td>(.194)</td>
</tr>
<tr>
<td>( \beta_6 )</td>
<td>.200</td>
</tr>
<tr>
<td>( \beta_7 )</td>
<td>(.393)</td>
</tr>
<tr>
<td>( \beta_8 )</td>
<td>.15</td>
</tr>
<tr>
<td>( \beta_9 )</td>
<td>.48</td>
</tr>
<tr>
<td>( \beta_{10} )</td>
<td>1.32</td>
</tr>
</tbody>
</table>
a term). The figures in parentheses are the (conditional) standard errors of the coefficients which appear above them; these were not computed for the coattails parameter (row (6)), and also not for the intercept term (row (1)), because of the way the regression program was set up. The last two rows contain the values of the squared multiple correlation coefficient and the Durbin-Watson statistic from the regression corresponding to $\gamma = \gamma'$.

All of the equations explain a significant amount of the vote: if we apply a likelihood ratio test of the hypothesis that all coefficients are zero except the intercept term and the coefficient of $T$, each of the equations is significant at better than .01. The equations including a trend account for almost two-thirds of the variance of the vote (after eliminating the estimated effect of the presidencial disturbance $e$), while the others explain about half the variance.

With respect to the individual estimates, the estimated coefficients of the income terms (both monetary and real) are positive in every equation, as expected a priori; moreover, judging from their conditional standard errors, they are also highly significant. The price coefficients also have the proper sign (negative) in every case, but are significant only in the equations (3.3) and (3.6)) which involve monetary income $Y$. In those equations, fluctuations in prices and monetary incomes have effects which are similar in magnitude, though opposite in sign; when real-income fluctuations are considered, however, price fluctuations apparently have little independent effect. The unemployment coefficient consistently appears with the wrong sign (positive), but it also invariably has a very large standard error, so is never significant. Two of our a priori coefficient expectations are therefore confirmed, and the third is at least not refuted.

The fact the unemployment fluctuations have no significant effect is somewhat puzzling. Rees et al. found some association between insured unemployment and the Republican vote,\(^{21}\) and Durant cites several investigations in Britain which found correlations between unemployment and the popularity of the Government during the past decade.\(^{22}\) However, none of these studies attempted to simultaneously consider the effects of price or income fluctuations in a systematic fashion, so it is difficult to compare their results with ours. As mentioned earlier, alternative forms of the unemployment term were also tried in the various equations; although the corresponding coefficient was sometimes negative (as expected a priori), in every case the standard errors were large, and no consistently significant effect could be found. The fact that the unemployment term is somewhat intercorrelated with the income term ($r = - .77$, with income in current dollars) may be part of the problem. Errors in the unemployment series, especially for the pre-World War II period, undoubtedly also play a role. A third factor may be the fact that, at normal unemployment levels, the unemployed consist disproportionately of those types of individuals who tend to be least active politically (often because of racial discrimination or inability to meet residency or poll tax qualifications), and hence may have relatively little direct impact on aggregate vote statistics. In any event, for whatever reason, our results show no significant independent unemployment effect.

The coefficient of the incumbency variable $\delta$ is invariably small and insignificant. In view of the substantial income effects in every equation, the most reasonable interpretation is that the institutional advantages of incumbency are approximately offset by high income expectations on the part of the electorate. The trend term $T$ is clearly significant, and important; the underlying partisanship of the electorate was about 54% Republican in 1896, and changed progressively to about 47% at the end of the period. The importance of the trend is also reflected in the Durbin-Watson $d$ statistic, reported in row (10) of Table 1. As noted earlier, the usual critical regions for this statistic are not strictly applicable here (and in any event, all values lie in the inconclusive region); however the statistic is consistently closer to 2 in the equations containing the $T$ term, reflecting a lessening of the apparent serial correlation of the residuals.

In the equations which do not contain the trend term, the estimated coattails parameter is zero (based on a .01 grid). Because of the importance of the trend, the estimates in (3.4), (3.5) and (3.6) are more relevant; there, the estimated value is $\hat{\gamma} = .15$ (based on a .05 grid). If we apply a likelihood ratio test of the hypothesis that $\gamma = 0$, though, then in every case the estimate is not significantly different from zero, so our evidence still does not indicate any important coattails effect. However, the manner in which the minor-party vote was treated in obtaining these results has probably biased these estimates downwards. Nationally based minor parties which run presidential candidates usually do not enter candidates in most House races, so the possible coattails of their presidential candidates have little opportunity to manifest themselves; and conversely, in the congressional vote totals, a large part of the minor-party vote comes from independent candidacies, or from state-based minor parties with no national organizations (such as the Liberal party of New York), so that in these cases the possibility of presidential coattails does not even arise. For these (and possibly other) reasons, increases in the minor-party vote have usually been associated

---

\(^{21}\) Rees et al., op. cit.
\(^{22}\) Durant, op. cit.
with increases in split-ticket voting. Hence we may get a better idea of the magnitude of presidential coattails by deleting the minor-party vote altogether, and considering only the two-party vote totals for congress and president. If the six equations referred to above are reestimated with the dependent variable redefined in this fashion, in every case we find that \( \gamma = .3 \) (based on a .05 grid), and the estimates are now all significantly greater than zero. This implies that around one-third of the votes gained (or lost) because of the specific candidates and campaign tactics of the presidential race carry over to the congressional candidates of the same party. The total variance of the major-party vote (based on equation (3.5), as reestimated here) can be decomposed as follows:

<table>
<thead>
<tr>
<th>Table 2. Variance of the Major-Party Vote</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Source</strong></td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td>( 2\delta X )</td>
</tr>
<tr>
<td>( u )</td>
</tr>
<tr>
<td>( v )</td>
</tr>
<tr>
<td>( \gamma )</td>
</tr>
</tbody>
</table>

In congressional elections held in presidential years, presidential coattails account for about one-sixth of the variance of votes, and economic variables for close to half. In presidential elections, by contrast, less than a fifth of the variance is due to these economic factors, while the presidential disturbance \( \eta \), which represents the effects of such factors as the personalities and campaigning abilities of the candidates themselves, account for two-thirds of the variance. Presidential races are thus substantially less predictable than congressional elections, as expected.

Further insight into the overall performance of the models can be obtained by examination of their predictive ability. In Figure 1 I have plotted the predicted "team" vote (the solid line), based on equation (3.5), and also the actual Republican presidential and congressional votes (circles and triangles, respectively), for each election. (The predictions based on the other equations containing the trend term are quite similar.)

The actual congressional vote is generally within a few percentage points of the predicted vote, except for the 1920s and a few individual years, such as 1898. As noted earlier, the predictions account for about two-thirds of the total variance. Another test of the forecasting performance of a model, frequently used in econometric work, is its ability to predict turning points, or changes in trend. The congressional vote has shown no strong trend over the period considered, and some twenty-two of our elections constitute turning points according to the usual definition, so this test is not useful. However, if we simply consider its ability to predict increases or decreases in the Republican vote share, the model performs reasonably well: if a forecast increase is defined as the difference between the current predicted vote and the actual vote in the preceding election (or in the last election used for estimation), twenty-five of the thirty forecasts are correct in sign. If the difference between current and previous predicted vote is used as the forecast, the sign of the forecast is still correct twenty-four times.

During most of the 1920s, the actual Republican vote was consistently larger than that predicted by the model. The composition of the U.S. electorate changed substantially during this period, as a result of the enfranchisement of women in 1920, and our simple model does not take account of this important factor. With respect to other years, there are substantial residuals in 1898, 1932, 1950 and 1964. In 1898 and in 1950, U.S. troops were involved in hostilities abroad, and in both years the vote for the incumbent was appreciably less than predicted, which is suggestive. (The residuals in the immediate post-war elections of 1920 and 1946, though smaller in magnitude, also follow the same pattern.) The model underestimated the magnitudes of the Democratic landslides of 1932 and 1964, which is not surprising in view of the special circumstances of those elections. Overall, the congressional vote does not deviate from the predicted values in a systematic or inexplicable fashion.

**IV. Summary and Conclusions**

One basic finding to emerge from this study is that election outcomes are in substantial part responsive to objective changes occurring under the incumbent party; they are not "irrational," or random, or solely the product of past loyalties and habits, or of campaign rhetoric and merchandising. In this respect, our findings support those of Key, based on quite different data.

Economic fluctuations, in particular, are important influences on congressional elections, with economic upturn helping the congressional candidates of the incumbent party, and economic...
decline benefiting the opposition. In quantitative terms, a 10% decrease in per capita real personal income would cost the incumbent administration 4 or 5 percent of the congressional vote, other things being equal. With a "swing ratio" of 2, this would translate into a loss of around 40 House seats—a respectable shift. In analysis-of-variance terms, economic fluctuations can account for something like half the variance of the congressional vote, over the period considered. (Presidential elections, however, are substantially less responsive to economic conditions.)

Of the economic variables considered, real personal income seems to be the most important; with real income held constant, changes in unemployment or in the rate of inflation have no significant independent effects. Being of the same party as the incumbent administration is of benefit only to the extent that the administration is successful economically; there is little intrinsic benefit to this type of incumbency in itself. Presidential coattails, however, are important, with perhaps as much as 30% of the extra votes attracted by a strong presidential candidate, who runs well ahead of his ticket, benefiting the congressional candidates of his party.

Appendix A

This discussion of estimation is necessarily somewhat technical. Good introductory expositions to the principles involved can be found in the books by Freeman, and Mood and Graybill; references to the pertinent pages are given below.

Let $y_i^t$ and $y_j^t$ be the Republican shares of the presidential and congressional votes in election $t$, observed over $T_1$ separate presidential elections, and let $X_{1i}^t$, $\ldots$, $X_{ki}^t$, $u^t$, and $v^t$, be the values of the $k$ independent and two random variables in that year $t$. Then we have

$$
    y_i^t = \beta_1 X_{1i}^t + \beta_2 X_{2i}^t + \cdots \\
    + \beta_k X_{ki}^t + u^t + v^t
$$

$$
    y_j^t = \beta_1 X_{1j}^t + \cdots \\
    + \beta_k X_{kj}^t + u^t + v^t
$$

$$
    t = 1, 2, \ldots, T_1
$$

For each of the $T_2$ off-year elections, let $y_{ij}^{t''}$ be the congressional vote, and $Z_{1i}^{t''}$, $\ldots$, $Z_{ki}^{t''}$, and $w^{t''}$ be the explanatory and random variables, so that

$$
    y_{ij}^{t''} = \beta_1 Z_{1i}^{t''} + \cdots + \beta_k Z_{ki}^{t''} + w^{t''},
$$

$$
    t'' = 1, 2, \ldots, T_2
$$

The disturbances are distributed normally and independently of one another, and their moments are assumed to satisfy.
\[ E(u^{t'}) = E(v^{t'}) = E(w^{t'}) = 0, \quad \text{all } t \]
\[ E(u^{t}w^{t'}) = E(w^{t}v^{t'}) = \begin{cases} 0, & t \neq t' \\ \sigma_u^2, & t = t' \end{cases} \quad (A.2) \]
\[ E(v^{t'}v^{t''}) = \begin{cases} 0, & t \neq t' \\ \sigma_v^2, & t = t' \end{cases} \]
\[ E(u^{t}v^{t'}) = E(w^{t}w^{t'}) = E(v^{t}w^{t'}) = 0, \quad \text{all } t, t' \]

In view of these assumptions the conditional joint density of the dependent variables can be written
\[ f(y_1^{t}, \ldots, y_N^{T_1} | X_1^{t}, \ldots, X_k^{T_2}) \]
\[ = \prod_{t=1}^{T_1} f_1(y_1^{t}, y_2^{t} | X_1^{t}, \ldots, X_k^{t}) \cdot \prod_{t'=1}^{T_2} f_2(y_2^{t'}, Z_1^{t'}, \ldots, Z_k^{t'}) \quad (A.3) \]

To obtain the likelihood function we first note that
\[ f_1(y_1^{t}, y_2^{t} | X^{t'}) \]
\[ = f(u^{t}, v^{t} | X^{t'}) \left| \frac{\partial (u, v)}{\partial (y_1, y_2)} \right| + \]
\[ = f(u^{t})f(v^{t}) \frac{\partial (y_1, y_2)}{\partial (u, v)} \bigg|^{-1} \bigg| \left. \frac{\partial (u, v)}{\partial (y_1, y_2)} \right| + \]
\[ = \frac{1}{\sqrt{2\pi \sigma_u}} \frac{1}{\sqrt{2\pi \sigma_v}} \exp \left[ -\frac{1}{2} \frac{u_1^2}{\sigma_u^2} \right] \cdot \frac{1}{\sqrt{2\pi \sigma_v}} \frac{1}{\sqrt{2\pi \sigma_v}} \exp \left[ -\frac{1}{2} \frac{v_1^2}{\sigma_v^2} \right] \bigg| \left. \frac{1}{\gamma} \right|^{-1} \bigg| \left. \frac{1}{\gamma} \right| + \]
\[ = \frac{1}{2\pi \sigma_u \sigma_v (1 - \gamma)} \exp \left[ -\frac{1}{2} \left( \frac{(u')^2}{\sigma_u^2} + \frac{(v')^2}{\sigma_v^2} \right) \right]. \quad (A.4a) \]

Similarly,
\[ f_2(y_2^{t'} | Z^{t''}) \]
\[ = f(w^{t'} | Z^{t'}) \left| \frac{\partial w}{\partial v} \right|^{-1} \bigg| \left. \frac{\partial w}{\partial w} \right| + \]
\[ = \frac{1}{\sqrt{2\pi \sigma_u}} \frac{1}{\sqrt{2\pi \sigma_v}} \exp \left[ -\frac{1}{2} \frac{(w')^2}{\sigma_u^2} \right]. \quad (A.4b) \]

Hence, substituting into (A.3), the density becomes
\[ f(y | X, Z) = 2\pi^{-(T_1+T_2)/2} \sigma_u^{-T_1} \sigma_v^{-T_2} (1 - \gamma)^{-T_1} \]
\[ \cdot \exp \left[ -\frac{1}{2} \left( \sum_{t=1}^{T_1} \frac{u_1^2}{\sigma_u^2} + \frac{v_1^2}{\sigma_v^2} \right) \right] + \sum_{t'=1}^{T_2} \frac{w_{1t'}^2}{\sigma_u^2} \right] \]

To express this in terms of the observed variables \( y, X, Z \), we first solve (A.1) to obtain
\[ u^{t'} = q^{t'} - (\beta_1 X_1^{t'} + \beta_2 X_2^{t'} + \cdots + \beta_k X_k^{t'}) \]
\[ v^{t'} = \frac{1}{1 - \gamma} (y_1^{t'} - y_2^{t'}) \quad (A.6a, b, c) \]
\[ w^{t'} = y_2^{t'} - (\beta_1 Z_1^{t'} + \cdots + \beta_k Z_k^{t'}) \]
where
\[ q^{t'} = -\frac{\gamma}{1 - \gamma} y_1^{t'} + \frac{1}{1 - \gamma} y_2^{t'} \]
is a transformation of \( y_1^{t'} \) and \( y_2^{t'} \) which depends on the unknown structural parameter \( \gamma \). If we substitute (A.6) into (A.5), the density is expressed entirely in terms of the observed variables and structural parameters; if we then regard the observations as fixed and consider the expression to be a function of the parameters, we obtain the logarithmic likelihood function\(^7\)
\[ L = -\left( T_1 + \frac{T_2}{2} \right) \ln 2\pi - (T_1 + T_2) \ln \sigma_u \]
\[ - T_2 \ln \sigma_v - T_1 \ln (1 - \gamma) - \frac{1}{2} \sum_{t=1}^{T_1} \left( q - \beta_1 X_1^{t} - \cdots - \beta_k X_k^{t} \right)^2 \sigma_u^2 \]
\[ - \frac{1}{2} \sum_{t'=1}^{T_2} \left( y_1^{t'} - y_2^{t'} \right)^2 \sigma_v^2 \]
\[ - \frac{1}{2} \sum_{t'=1}^{T_2} \left( y_2^{t'} - \beta_1 Z_1^{t'} - \cdots - \beta_k Z_k^{t'} \right)^2 \sigma_u^2 \]

The joint maximum-likelihood estimates are those values of the parameters \( \beta_1, \ldots, \beta_k, \gamma, \sigma_u^2, \sigma_v^2 \) which maximize the likelihood. If we attempt to simply apply the usual first-order conditions for a maximum to (A.8), the resulting expressions will be awfully nonlinear in the coattails parameter \( \gamma \), and difficult to solve numerically. However, the maximum can be approximated as closely as \( ^7 \) See Freeman, op. cit., pp. 253c; and Mood and Graybill, op. cit. pp. 178ff.
desired by a simple stepwise procedure, as follows. If the coattails parameter \( \gamma \) were known, or if a value could be assumed for it, maximum likelihood estimates for the remaining parameters are easily obtained. As is readily verified, the resulting estimates \( \hat{\beta} \) of the \( \beta \) are the values that would be obtained from an ordinary least-squares multiple regression with dependent and independent variables being respectively \( q^t \) and \( X^t, \ldots, X^t \) for the \( T^t \) observations of presidential years, and \( y^t, Z^t, \ldots, Z^t \) for the remaining \( T^t \) observations. The estimate \( \hat{\sigma}^2 \) of the variance of the "team" disturbance \( u \) is \( \Sigma e^t/(T^t + T^t) \), where \( \Sigma e^t \) denotes the residual sum of squares from the above regression, and the estimate \( \hat{\sigma}^2 \) is \( 2\Sigma (y^t - y^t)/T(1 - \gamma) \). These estimates are conditional on \( \gamma \), and as different values of \( \gamma \) are assumed, different estimates will be obtained. If the expressions for \( \hat{\beta}, \hat{\sigma}, \hat{\sigma}^2 \) are substituted for the corresponding unknown parameters in (A.8), we obtain an expression for the conditional maximum of the likelihood \( \mathcal{L}(\gamma) \) of the form \( \mathcal{L}(\gamma) = K - (T^t + T^t) \Sigma e^t/2 \), where \( K \) is a term not involving any of the parameters; the likelihood is therefore at a maximum when the residual sum of squares \( \Sigma e^t \) is minimum. Our estimating procedure therefore consists of computing the regression and sum of squares using each of a succession of trial values for \( \gamma \), e.g. 0, .05, .10, .15, etc., until the value \( \gamma \) which results in the smallest residual sum of squares has been (approximately) found; since it is known that \( \gamma \) lies in the interval \( 0 \leq \gamma < 1.0 \), this is not difficult computationally. The resulting \( \gamma \) and associated \( \hat{\beta}, \hat{\sigma}, \hat{\sigma}^2 \) are the (unconditional) joint maximum-likelihood estimates of the parameters of the model.

Likelihood ratio tests\(^a\) are readily constructed to test hypotheses on the parameters. The likelihood ratio for the hypothesis that the parameters satisfy \( r \) independent linear constraints (such as \( \beta_1 = 0, \ldots, \beta_0 = 0 \)) can be shown to be \( -(T^t + T^t) \Sigma e^t - \ln \Sigma e^t \), where \( \Sigma e^t \) is the residual sum of squares from the regression with the constraints imposed, and \( \Sigma e^t \) is defined as before; for large samples this test statistic will be distributed approximately as a chi-square with \( r \) degrees of freedom.

Appendix B

The primary data sources used were Historical Sta-

\(^a\) See Freeman, op. cit., pp. 300ff; and Mood and Graybill, op. cit., pp. 297ff.

suchistics of the United States (1957), and Statistical Abstract of the United States, various issues. These two sources will be referred to as HSUS and SAUS, respectively, in this appendix.

Population. Total population for 1895–1940 from HSUS, p. 7 (values prior to 1899 obtained by linear interpolation between 1889 and 1899 figures); for 1941–1964, SAUS (1966), p. 616.


Unemployment. As percent of civilian labor force.

"Prior to 1940, these figures represent estimates of unemployment on as comparable a basis to current labor force concepts as is presently possible. . . . Unemployment is calculated as a residual. That is, estimates are first made of the civilian labor force, then of employment; the difference between the two provides the estimates of unemployment." (HSUS, p. 68.)

After 1940, the unemployment figures are based on field surveys. The figures for 1900–1939 were taken from HSUS, p. 73 (prior to 1900, set equal to the 1900–03 average), for 1946–61, from SAUS (1962), p. 215, and for 1962–64, from SAUS (1966), p. 218.


"This index . . . measures the average change in prices of goods and services purchased by city wage-earner and clerical-worker families. . . . The weights used in calculating the index are based on studies of actual expenditures by wage earners and workers." (SAUS (1966), pp. 349–50, cf. also HSUS, pp. 109–10)

For 1895–1912, the Federal Reserve Bank of New York cost-of-living index was used. This index was obtained by splicing together parts of indices already available to approximate a single series. No adjustments were made to the original series other than those necessary to convert to a common base period. . . . For 1890–1909, Paul Douglas' "most Probable Index of the Total Cost of Living for Workingmen," was used. Indexes for 1910–12 were derived from the cost-of-living index for Massachusetts. . . ." (HSUS, p. 111)

Figures were obtained from HSUS, p. 127, and adjusted to the same base period as the BLS series described above.

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