The Electoral Information Hypothesis Revisited*

Jude Hays
University of Michigan

Helmut Stix
University of Vienna and
Osterreichische Nationalbank

John R. Freeman
University of Minnesota

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Abstract

The connection between political and bond market equilibration is studied. The conventional wisdom on the subject, the Electoral Information Hypothesis (EIH; Cohen 1993, Alesina, Roubini and Cohen 1997), is criticized on theoretical and empirical grounds. Using a framework proposed by Garrett and Lange (1995) and data from three of the most developed bond markets in the world a revised hypothesis is constructed and tested. The revised hypothesis recognizes the way formal political and bureaucratic institutions mitigate the effects of democratic politics on bond markets as well as the "fat tailed" distributions and volatility clustering commonly found in financial time series. The results show that empirical support for the revised EIH extends beyond the American case, but this support is not universal. More specifically, where democracy is of the majoritarian type and central banks are weak (the U.K. until recently) the revised EIH holds. Once placed on sounder statistical footings, the U.S. case—one of mixed majoritarian and consensual democracy and a moderately strong central bank—also supports the revised EIH. However, the revised EIH does not receive support in Germany, which has a consensual form of democracy and a strong central bank. The empirical power of the revised EIH thus is shown to vary by institutional context.

The study of democracy and markets is at the heart of political economy. Understanding how political equilibration and economic equilibration are related is one of the main challenges facing this field. The size and nature of bond markets make them especially important cases. Substantively, with globalization, bond markets have become important constraints on elected governments, reducing their "room to maneuver." In fact, there is increasing concern among political economists about the constraints that the globalization of bond and other financial markets impose on democracy. Analytically, in comparison to goods markets, bond markets are informationally efficient asset markets with distinctive equilibria. Bond prices presumably reflect traders' fully informed expectations about, among other things, future rates of inflation. The sensitivity of these markets to the arrival of new information therefore makes them especially well suited for studying the economic implications of political news and uncertainty. The connections between political and economic equilibration should be evident in the behavior of bond markets.[1]

The Electoral Information Hypothesis (EIH) is representative of the conventional wisdom about political-economic equilibration in general and political and bond market equilibration in particular.[2] The EIH is based on two simple ideas: (1) political parties adopt distinctive policies and hence produce different macroeconomic outcomes—especially distinctive patterns of inflation, and (2) because inflation affects investors' real rates of return, forward looking-financial markets react immediately to information about the likelihood of changes in the partisan composition of government. The result is a smooth

[1] For a review of the literature on political economic equilibration in financial markets see Cohen (l993, 17-18), Alesina, Roubini, and Cohen (l997, 126-8), Mosley (1999), and Herron (2000). The differences between goods and financial markets are explored in such works as Hallwood and MacDonald (l994).

[2] Other important recent works on this subject are Mosley (1999) and Perry and Robertson (1998). The links between these studies and the EIH are explained below.

transition of these markets across electoral equilibria or, few post-electoral discontinuities in bond market behavior. Cohen (l993) and Alesina, Roubini and Cohen (l997) formulate this hypothesis in terms of the concept of forward rate revision of bonds. They show that in the United States for bonds of varying maturities and revisions of one and three months, there is much evidence that changes in the probability of the election of more inflation prone Democratic governments affect bond yields. Increases in these probabilities cause upward shifts and/or a widening in the term structure of interest rates. Since the U.S. is among the countries with the greatest degree of financial openness, this finding suggests that both domestic and foreign traders are sensitive to political information; globalization is consistent with a sequence of smooth electoral-bond market equilibria.[3]

This paper critically evaluates the EIH. It argues that the EIH is theoretically underdeveloped and empirically unsound. The hypothesis is underdeveloped theoretically in that it does not make any provision for the workings of non-American political institutions. For instance, the fact that many political systems do not provide at the executive level for winner-take-all election outcomes but rather for the formation of coalitions of parties is not incorporated in the EIH. Nor does the EIH take into account bureaucratic constraints on executives such as central bank independence. Empirically, the reduced form equation that Alesina, Roubini and Cohen employ does not account for the excess kurtosis and serial correlation in the conditional second moments of financial time series, substantively, the fact that these series exhibit extreme values surprisingly often and that these extreme values tend to

[3] Because a bond entitles the owner to a fixed income payment (or stream of payments) there is a negative relationship between a bond's price and its yield. Hence when the yield curve is shifting upward, bond prices are decreasing. This happens when traders have inflationary expectations. As regards the financial openness of the U.S., in the l980s the level of public debt held by foreigners averaged $222.6 billion; in the l990s this average was $741 Billion. U.S. public debt held by foreigners as a proportion

cluster together in time.[4] Current economic thinking holds that univariate interest rate and term structure dynamics are governed by regime-switching processes.

A revised hypothesis about political and bond market equilibration is presented, one that incorporates many of the features of the EIH. The new hypothesis explains why politics does not affect bond markets in some democracies. We translate the new hypothesis into a model that provides for interest rate regime switching along the lines proposed by Hamilton (l988), Ang and Bekaert (l998) and others.[5] We show that this model performs well for forward rate revisions in a number of advanced industrial countries, countries with a wide variety of institutional configurations. The model shows support for the EIH is not universal, however. Where democracy is of the majoritarian type and central banks are weak (the U.K. until recently) the revised EIH holds. Once placed on sounder statistical footings, the U.S. case—one of mixed majoritarian and consensual democracy and a moderately strong central bank—also supports the revised EIH. However, the revised EIH does not receive support in Germany, which has a consensual form of democracy and a strong central bank. The empirical power of the revised EIH thus is shown to vary by institutional context.[6]

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of all non-bank investment in U.S. public debt issues averaged .212 in the l980s and .267 in the l990s. Source: *Economic Report to the President*, 1998.

[4] Excess kurtosis or "fat tails" (the relatively frequent occurrence of extreme values) and serial correlation in the conditional second moments (autoregressive conditional heteroscedasticity) are common properties of forward rate revisions and other financial time series. Even the International Monetary Fund (1998, 6-9) now makes reference to these "fat tailed events" in financial markets. For example, the IMF explains the failure of value at risk models in the recent Asian crisis in terms of an inability to predict losses from "fat-tailed events" and associated changes in correlations and volatilities across markets. See also Kim and Nelson, 1999.

[5] For example, Gray (1996), Garcia and Perron (1996), Evans (1998), and Bansal and Zhou (2000).

[6] These results build on Freeman, Hays, and Stix (2000) who find that plurality electoral systems exacerbate the impact of politics on currency market equilibration.

The discussion is divided into four parts. Part one critically evaluates the EIH. The revised EIH is developed in part two. A research design for evaluating it is presented and executed in part three. The challenge and importance of understanding political-economic equilibration in financially open systems are discussed briefly in the conclusion.

The EIH Revisited

The EIH argues that in the period running up to an election, asset prices reflect traders' assessments of the electoral prospects of competing parties. Consequently, the actual outcomes of electoral contests have little effect on financial markets; these outcomes already have been anticipated and incorporated in prices. Insofar as financial markets are concerned then, the transition between political-economic equilibria usually is quite smooth. This smooth transition between electoral equilibria is the key stylized fact motivating the EIH.[7]

As applied to the bond market in the U.S., the EIH contends that electoral information is reflected in the prices and hence yields of Treasury bills. More specifically, as the probability of a Democratic (Republican) presidential victory increases, bond traders expect higher inflation and hence long-term bond yields and forward interest rates increase (decrease). Cohen (l993) and Alesina, Roubini, and Cohen (l997) develop this argument in terms of the expectations theory of the term structure of interest rates. This theory holds that the time t interest rate on a k-period bond will be equal to the sum of the current one-period interest rate and the expected interest rates on the k-1 one-period bonds that span the same investment horizon from time t to time t+k. Forward interest rates thus represent expected future interest rates and consequently are a function of expected real interest rates,

[7] The transition between equilibria is more discontinuous when traders are surprised by the electoral outcome as in l948 and perhaps l992. See Cohen (l993).

expected inflation, and a liquidity premium. On the basis of rational partisan theory, Alesina, Roubini and Cohen argue that expected inflation is determined by the expected probability of a Democratic presidential victory and the inflation differential associated with that executive's policies. In this way, electoral information is a determinant of forward interest rates. This is the essence of the EIH.[8]

The key relationship on which Alesina, Roubini and Cohen focus is

$$\Delta_d Fk_{j-d,t} = (r^e_{t+j-d,k|t} - r^e_{t+j-d,k|t-d}) + (P^D_{t+j-d|t} - P^D_{t+j-d|t-d})(p^D - p^R) + \Lambda_{k,t}\ , \tag{1}$$

where $\Delta_d Fk_{j-d,t}$ is the forward rate revision from time t-d to time t for a bond of maturity k purchased at time $t+j$-d, $r^e_{t+j-d,k}$ is the expected real return on a k-maturity bond purchased at time $t+j$-d, P^D_{t+j-d} is the subjective probability of the Democrats winning the next presidential election and therefore the probability of their policies being in place at time $t+j$-d, p^D and p^R are the inflation rates under Democratic and Republican executives respectively, and $\Lambda_{k,t}$ is the d-period change in the term or "liquidity" premium at time t. Conceptually, the forward rate revision is the change in expectations from time t-d to t of the rate on the same k-period bond purchased at time $t+j$-d.[9,10] The term premium

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[8] In comparison to other financial assets, Treasury bills presumably are more likely to reflect national electoral outcomes like changes in the identity of the party controlling the Presidency than stocks which are more affected by the outcome of particular Congressional elections (Cohen l993, 128-9). See Mosley (1999, Chapter 1) for the importance bond traders attach to inflation expectations. Perry and Robertson (1998, 134) also stress that it is the "potential for policy shifts" especially with respect to inflation—that define the risk bond traders face. We return to the connection between the EIH and work on alternative interest rate instruments and security prices including Herron (2000) in our Discussion section.

[9] Forward rates thus are a key element of the larger theory that relates current prices to expected rates of return. For an introduction to the concept of forward rate revision see our Appendix and Appendices B and C of Alesina, Roubini and Cohen (1997). An introduction to forward interest rates and to the expectations theory of the term structure can be found in such works as Fabozzi, Modigliani, and Ferri (l994, Chapters 11 and 12).

[10] In the appendix we connect the concept of forward rate revision to the literature on term structure estimation. In general the notation used in this literature is unavoidably messy. To simplify it we use a different notation than Alesina, Roubini and Cohen.

denotes, among other things, a liquidity preference for bonds of a certain maturity. On the basis of some work in economics, Alesina, Roubini, and Cohen argue that the term premium is either constant or that stochastic variation in this term is small relative to the variation in the other right hand side variables. Hence they ignore $\Lambda_{k,t}$ in their analyses.

Alesina, Roubini, and Cohen employ the following reduced form

$$\begin{aligned}\Delta_d Fk_{j-d,t} = \boldsymbol{b}_0 + \boldsymbol{b}_1 \Delta_d Fk_{j-d,t-d} + \boldsymbol{b}_2 \Delta_d Fk_{j-d,t-(d+1)} + \boldsymbol{b}_3 \Delta_d \boldsymbol{p}_{t-1} + \\ \boldsymbol{b}_4 \Delta_d U_{t-1} + \boldsymbol{b}_5 \Delta_d M1_{t-1} + \boldsymbol{b}_6 \Delta_d P_t^D + \boldsymbol{e}_t\end{aligned},$$

(2)

where $\Delta_d \boldsymbol{p}_{t-1}$ is the d-month change in the inflation rate at time *t-1*, $\Delta_d U_{t-1}$ is the d-month change in the unemployment rate for time *t-1*, $\Delta_d M1_{t-1}$ is the d-month log change in M1 at time *t-1*, and $\Delta_d P_t^D$ is the d-month change in the probability of a Democratic President being elected as calculated at time t. The first and second terms on the right hand side of (2) are intended to correct for serial correlation in $\boldsymbol{e}_t$. The third, fourth, and fifth terms are proxies for real interest rate changes.

Alesina, Roubini, and Cohen fit (2) for bonds of various revisions and maturities. Their results support the EIH. The $\boldsymbol{b}_6$ coefficients are positive and statistically significant, suggesting that in the U.S. changes in pre-electoral information cause an upward shift in and/or a widening of the term structure of interest rates. In particular, Alesina, Roubini, and Cohen (1997) find that a ten percent change in the calculated probability of a Democratic presidential victory (the mean absolute monthly change in their data set) causes an immediate 5-8 basis point change in the implied rate of 1, 2, 3, and 4-year bonds that are purchased eleven months in the future. Using simulation methods, the authors estimate that the magnitude of change witnessed in the pre-l992 electoral period—10 to 80% change in P_t^D — produces a 35-60 basis point increase in this forward rate for these bonds.[11]

Critique. The EIH suffers from several theoretical and statistical problems. To begin with, the EIH really is a theory of one type of polity, namely, one in which presidential (executive) elections have a fixed cycle and are of the winner-take-all variety. If the electoral system produces governing coalitions of parties, the nature of the hypothesis is much less clear. Take a proportional representation electoral system. Do bond and other traders focus on the pre-electoral fortunes of the ruling coalition (the expected votes of the respective parties combined) or, those of the leader of some "strong party"?[12] Also, central banks often enjoy tremendous autonomy over monetary policy: central banks often are "insulated" from the vagaries of electoral and legislative politics. Alesina, Roubini, and Cohen do not address, let alone answer, these questions. As we explain below, political theory suggests that bureaucratic and formal political institutions reduce, if not eliminate, the effects of elections and cabinet dissolutions on price trends. Political theory thus predicts that the EIH will not apply with equal force in all democracies. In addition, Alesina, Roubini and Cohen's reduced form (2) is problematic. The two lags on the right side of equation (2) suggest that traders react sluggishly to new political and economic information. But this is inconsistent with rational expectations assumption on which the EIH is based.

Empirically, the EIH is unsound. Alesina, Roubini, and Cohen's linear regression model does not account for the excess kurtosis and serial correlation in the conditional second moments of their financial time series. To demonstrate these deficiencies we replicate Cohen's (1993) results for three-

[11] A basis point is 1/100 of a percent. So an impact of 35-60 basis points is equivalent to a change of .35 - .60 percentage points in the forward rate revision.

month revisions in the forward rate of three-month bonds purchased three months ahead $\Delta_3 F3_{6,t}$. The results are reported in Table 1.[13] Analysis of the residuals from our replication shows that Cohen's model does not remove the excess kurtosis in the bond data. The distribution of the residuals has a mean of -.002 and a standard deviation of 1.035. There is excess mass in the tails of this distribution; instead of a value of 3 (normal distribution), kurtosis is 7.351 (Table 2). It is not surprising then that the hypothesis that the residuals are normally distributed is easily rejected (Jarque-Bera Statistic = 366.776, Table 2). This problem is shown graphically in Figure 1 where the probability density function for a N(-.002, 1.035) is plotted against the empirical distribution of the residuals. Put simply, we observe more extreme forward rate revisions in the sample than can be accounted for with Cohen's simple linear

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[12] Laver and Shepsle (l996) develop the concept of a "strong party." It is a party that because of its ideological stance is likely to be a member of all possible government coalitions.

[13] Cohen's (1993) research is the original work on which the EIH is based. We focus on Cohen (1993) rather than on the later work in Alesina, Roubini, and Cohen (1997) because the forward rate revisions in the former are more easily calculated with the data that is available in McCullough and Kwon (1993). Qualitatively, our results are virtually identical to Cohen's. The coefficients in our replication have the same signs and the coefficient on the political variable is statistically significant. The differences in our replication are due to three differences between Cohen's original analysis and ours: First, we used a slightly different sample. We were only able to obtain data back to 1948:5. Furthermore, we used the entire sample from 1948:5-1987:2 rather than omitting months for which Presidential trial-heat data is unavailable. Cohen's sample contains gaps (1993, 46) while ours does not. This is why our replication analysis is based on 466 observations while Cohen's analysis is based on 329 observations. We assume that bond traders only care about the election outcome if it occurs on or before *t+j-d*. The logic is that today's price for a three-month bond delivered in three months should not reflect the likelihood of a Democrat being elected President 2 years hence. Our political variable, $\Delta_d P^D_t$, takes a value of 0 for months that are deemed too early to have an impact on the bond market. Hence, we have no missing data. Second, we use a different measure of the money supply than Cohen. And finally, we use the Electoral Option Model to calculate the probability of a Democratic election victory rather than using the raw poll results. This accounts for the scale differences between the respective coefficients. Again, qualitatively, the results are very similar and the differences do not affect the basic points we are making: like most financial time series, forward rate revisions exhibit excess kurtosis and autoregressive conditional heteroscedasticity. A simple linear regression model like that employed by Cohen cannot account for these properties.

regression model.[14] Furthermore, Cohen's residuals display autoregressive conditional heteroscedasticity (ARCH). Using Ljung-Box tests, we can overwhelmingly reject the hypothesis that his squared residuals are not serially correlated (Table 2). This volatility clustering is clearly visible in a time series plot of Cohen's residuals (Figure 2). In these ways, the EIH rests on weak statistical footings. The distributional assumptions underlying Cohen's and Alesina, Roubini, and Cohen's hypothesis tests are problematic and their estimates in all likelihood are inefficient.

Finally, Alesina, Roubini and Cohen's analysis papers over seemingly important shifts in the character of political and bond market processes. In particular, there is much evidence that structural changes in these processes occurred in the late 1960s and again in the late 1970s (see, for instance, Perry and Robertson 1998). Also, many scholars have argued that differences in partisan preferences for inflation, which are assumed constant in Alesina, Roubini and Cohen's analysis, diminished in the 1980s. This raises the question of whether the EIH holds up in the most recent subperiod. In fact, if we split Cohen's sample into pre and post-1980 subsamples and reestimate his model, we find that the change in the expected probability of a Democratic election victory does not have a statistically significant impact on forward rate revisions for the later subperiod (Table 1, Column 4).

In sum, the EIH is a clear advance in our understanding of the connection between political and financial market equilibration. However, as we have shown, it is in need of revision.

A New and Improved EIH

[14] If the residuals in Figure 1 were generated by random draws from a normal distribution, the distribution most likely to have produced these draws has a mean of -.002 and a standard deviation of 1.035. We observe too many extreme values (i.e., too many near-zero probability draws) for this to be the case.

Let us first address the empirical shortcomings of the EIH. Excess kurtosis usually is attributed to regime switching, more specifically, the propensity of traders to switch between dynamic market equilibria or what are called market "regimes" (Hamilton, l988; Ang and Bekaert, l998). Operationally, switching between bond market equilibria amounts to probabilistic transitions between two models of forward rate revision, each with different parameters. Substantively the idea is that financial traders behave in different ways—adopting different (optimal) decision rules—depending on the information they receive; the sudden arrival of particular kinds of news causes traders to switch to alternate modes of behavior thereby creating different parameterizations of the same or sometimes different theoretical relationships. It is the mixing of these kinds of behavior that produces the excess kurtosis (fat tails) in the unconditional distribution of forward rate revisions and other financial time series.[15] Typically, there is a Markov process governing the switches between bond market regimes; there are probabilities at each point in time that traders will behave in the same way this period as they did in the previous period or switch to the alternate kind of behavior. These probabilities can be time varying.[16]

Our argument is that political information is, in part, responsible for the switches between bond market equilibria. The news bond traders receive about the condition of a prevailing political equilibrium causes them to alter their behavior in ways that produce discrete changes in the parameters of reduced form relationships for forward rates. In effect, the observation of (potential) political reequilibration

[15] See, for example, Hamilton's analysis of the densities of mixtures of two Gaussian distributions (1994, 685-8).

[16] Hamilton (l988, 394) discusses several alternative sources of such regime switching including the possibility that traders have nonlinear utility functions. He focuses on the idea that forecasts of future short-term rates are a nonlinear function of past short term rates. His final argument is that bond traders take changes in market regime into account and incorporate these changes in their forecasts of future forward rates according to the framework in the Rational Expectations Hypothesis of Term Structure.

causes traders to alter their optimal decision rules and, when aggregated, these changes produce a new parameterization of the forward rate equation. Operationally, pre-election polls and information about cabinet dissolution affect the transition probabilities between bond market equilibria. The greater the likelihood of the election or cabinet membership of an inflation prone party, the more likely the bond market will remain in or switch to a regime with an upward shifting (widening) term structure.

Figure 3 depicts the regime switching interpretation of the EIH. As we will show, this conception of the EIH is empirically sound. It explains the forward rate data—including the excess kurtosis and ARCH in forward rate revisions—much better than Alesina, Roubini, and Cohen's conception.[17]

The theoretical issue is whether political and bond market equilibration are related in this way in all political systems. Are the workings of political institutions such that pre-election polls and related kinds of political information cause switches between bond market equilibria in some countries but not in others? The implication of one body of work in political science is that the power of the EIH varies by institutional context. Illustrative is Garrett and Lange's (1995) argument that socioeconomic and formal political" institutions mitigate the effects of politico-economic change on public policy (Figure 4). Conceiving of elections as means of aggregating societal preferences about the Keynesian welfare state, Garrett and Lange argue that some institutions produce more stable, predictable policies that other

[17] Regime switching removes much but not all of the ARCH in the forward rate revision series. Therefore, we allow for conditional variances that change over time. Another reinterpretation of the EIH is that there are two or more parameterizations of (2) depending on the economic shocks that traders observe. And pre-electoral factors, $\Delta_d P_t^D$, only produce changes in forward rate revision in one of these regimes. It is economic shocks alone that determine which regime obtains. In our view, Alesina, Roubini, and Cohen's hypothesis is more consistent with the idea that political information causes traders to alter their optimizing behavior or that political uncertainty is part of the mechanism that causes regime switching.

institutions, policies that allow incumbents to protect their distributional interests in the face of shifts in the behavior of the electorate. Labor market organization is the key socioeconomic institution in this regard. The critical political institutions are constitutional factors—the workings of electoral rules in relation to the geographic location of incumbents' constituencies and the number of veto points in the system—and the power of certain bureaucracies.

For simplicity and because the best bond market data are available for democracies with distinct political institutions, we focus here on the impact of constitutional and bureaucratic factors on the EIH.[18] The constitutional factors Garrett and Lange analyze are elements of the somewhat broader distinction between majoritarian and consensus democracies (Lijphart 1999).[19] Majoritarian democracy downplays the need for unanimity at any point in time and equates popular sovereignty with majority rule; it usually is associated with plurality, single-member district electoral systems; few veto points are found in these systems. In contrast, consensus democracy emphasizes the need for mutual agreement among citizens and imposes constraints on majorities; it disperses and limits political power by, among other things, providing for multiple veto points. Proportional representation (PR) systems usually are found in consensus democracies. Garrett and Lange's argument suggests that the EIH is more likely to apply in majoritarian than consensual democracies. This is because the former type of democracy allows for comparatively quicker, more significant, and perhaps more unpredictable changes in economic policies. Majoritarian democracy' plurality electoral rules and single veto point means that

[18] Garrett and Lange's (1995) argument applies both to democratic and nondemocratic institutions. We focus in this paper only on democracies, and on the effect of formal political institutions. The respective countries are those for which we have the richest set of bond market data. We discuss the possibility of assessing the impact of the socioeconomic institutions on the EIH in our Conclusion.

the election of left parties is likely to result in changes in economic policies and, in turn, in changes in prices and other macroeconomic variables. It follows that in these democracies political information therefore should have an impact on the transition probabilities governing the swiches between bond market equilibria; for instance, pre-election polls that indicate a greater probability of the election of inflation prone left parties should increase the probability of remaining in or switching to a bond market equilibrium that connotes an upward shifting yield curve (see fn. 3). In consensual systems, in contrast, the power sharing produced by PR and gridlock created by multiple veto points, ought to produce few policy surprises and, more important, comparatively fewer significant changes in economic policy. Therefore, in this type of democracy, pre-election polls and information about cabinet formation and dissolution should not have an impact on the transition probabilities governing transitions between interest rate regimes; information about the likelihood of the election (government cabinet membership) of an inflation prone left party should not affect the probability of remaining in or shifting to a bond market equilibrium with an upward shifting yield curve in consensus democracies.[20]

The most important bureaucratic factor in relation to the EIH is the power of the central bank. Garrett and Lange and others argue that strong, independent central banks prevent changes in the partisan identity or coalitional makeup of government from affecting monetary policy and hence prices.

[19] Garrett and Lange (1995, 633) actually allude to the connection between their discussion of formal political institutions and Lijphart's work on democracies.

[20] As regards cabinet durability, the idea is that the proportional representation rules associated with consensus democracy consistently produce stable government coalitions. Illustrative is Laver and Shepsle's (1997) analysis of equilibrium cabinets. They show that the stability of these cabinets depends on such things as the cabinets being composed of strong parties—parties that because of the spatial distribution of preferences are always included in governments, there being a small number of parties, a small number of issue dimensions, and a decisive decision structure. When the equilibrium cabinet is a dimension-by-dimension median with an empty winset, its survival is resistant to many different kinds of political shocks. See also Rogowski (1988).

Where central banks are strong and independent, bond traders can safely ignore electoral politics and cabinet reorganizations; these events do not affect the course of monetary policy. Monetary policy remains in the hands of anti-inflationary central bankers. These bankers and private agents supposedly achieve reputational or rules equilibria that are immune from electoral and cabinet politics (Barro and Gordon 1983a,b).[21] Therefore, in countries with strong, independent monetary authorities, information about the electoral prospects of political parties and likelihood of cabinet reorganization should not have any impact on the probability of switches between bond market equilibria whereas the opposite will be true in countries with weak dependent monetary authorities.

In sum, this bond of political theory suggests that insofar as the impacts of "formal political are concerned (Garrett and Lange 1995), historically, the empirical power of the EIH will be greatest in majoritarian in majoritarian democracies with weak dependent central banks such as the United Kingdom and least in consensus democracies with strong independent central banks like Germany.[22] In the former kinds of democracies we should find the pre-election polls and information about the probability of cabinet reorganization affect the probability of switches between bond market equilibria; for example, increases in the probability of the election of left parties will increase the probability of remaining in or switching to an interest rate regime with an upward shifting yield curve.

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[21] Hall and Franzese (1998, 506-8) summarize this view. The crux of their argument is that central bank independence creates "credibility of assurance that monetary policy will remain tight, thereby allowing wage and price bargainers to lower their nominal contracts by reducing fears about real wage and real return losses that unanticipated inflation would create." Barro and Gordon (1983 a,b) study different equilibria that may arise between the monetary authority and private economic agents—equilibria with welfare consequences inferior to those produced by ideal (monetary policy) rules. Their suggestive reference to multiple equilibria of these kinds can be found in 1983a, section 10.

[22] Note the word "historically" here. We analyze the U.K. over the period 1980-1995. Recently the British central bank has been made more independent of elected offials.

But this will not be true in the latter kind of democracies; information about the electoral prospects of left parties and (or) cabinet reorganization will have no effect on the transition probabilities governing switching between bond market equilibria. For the reasons given at the start of this section, consensual systems with strong central banks may exhibit multiple, dynamic bond market equilibria—possibly corresponding to multiple reputational equilibria characterizing the bank's relationship with private agents (fn. 21). But, once more, the probabilities of shifts between these equilibria will not depend on political information.

Alesina, Roubini and Cohen's research bears on the intermediate case of the United States. The American political system has features of both majoritarian and consensual systems. For instance, its chief executive is elected in a winner-take-all plurality content while, overall, its legislative elections tend to produce outcomes that in some ways, resemble those of PR electoral rules (Garrett and Lange 1995, 644). The separation of powers in the American system provides for multiple veto points. Finally, the U.S. central bank is comparatively independent. So the fact that Alesina, Roubini, and Cohen found support for their original version of the EIH is important since it suggests that a high degree of consensual democracy and central bank in dependence must exist before bond markets are insulated from the effects of political equilibration. Of course, we do not know yet if Alesini, Roubini, and Cohen's results for the U.S. will hold up once the excess kurtosis and ARCH in forward rate revisions are addressed and the period since 1980 is studied. But, if the results do hold up, this would be the implication.

In fact, there are theoretical reasons to question whether even high degrees of consensual democracy and central bank independence insulate bond markets from electoral and other political events. Among democracies, the EIH could be universal. For one this, it is not clear that bond traders understand and appreciate the mitigating effects of institutions. For instance, traders may believe that left parties are able to subvert central bank independence. And there is no reason to believe that bond traders comprehend the process by which governments dissolve and form in consensual systems, let alone gauge the relative power of political parties within governing coalitions. As Laver and Shepsle (1996, 1997) have shown, this process is complex. Under some conditions, it can be stable and predictable. But under other conditions, government dissolution and formation can be chaotic.[23] For these reasons, the EIH could apply in consensual as well as majoritarian democracies. We thus have two competing propositions:

1. The power of the revised EIH depends on its institutional context. Switching between bond market equilibria depends on news about the prospects of inflation prone left parties coming to or remaining in power in majoritarian systems with weak central benks but not in consensual systems with strong central banks.

2. The power of the revised EIH is universal. News about the electoral prospects and cabinet membership of inflation prone parties affects bond market equilibration (switching) in all democracies.

Analysis

[23] Laver and Shepsle (1996) also show that under the opposite set of conditions (cf. fn. 15), cabinets can be quite unstable. In fact, when the party system has a dimension-by-dimension median with a nonempty winset, governments can cycle between different coalitions (1996, 68-9, 78ff). As regards sitting coalition governments, mid-term elections or poll results can change parties' expectations of future electoral outcomes and cause them to defect from government coalitions or to refuse to support votes of confidence for sitting governments. When these governments fall, new coalitions form with policy ideal points that can be significantly different than their predecessors. The conditions under which such events are observed have to do with such things as the proximity to the next election, existence of "very strong parties" in the government coalition and, of course, magnitude of the political shocks (Lupia and Strom 1995, Laver and Shepsle, 1996). The idea is that bond traders also make these inferences and adjust their decision rules accordingly. Hence we observe switches in bond market equilibria (regimes).

Design. To test Propositions 1 and 2, we need countries with different forms of democracy and whose central banks have different degrees of independence. In order to test our revised conception of the EIH we need countries whose bond markets are well developed; forward rate calculations are most meaningful where bond markets are "thick" and highly liquid. With these considerations in mind, we chose to (re)analyze the British, American, and German cases. The U.K. and Germany have majoritarian and consensual systems respectively; the U.S. system has features of both kinds of democracy. In our period of analysis, 1980:4-1995:12, the U.K. had a weak, dependent central bank whereas the central banks of the U.S. and Germany were comparatively strong and independent. The British, American, and German bond markets are among the most highly developed in the world.[24] By studying the effects of political information on the bond markets of these three countries we therefore can test our theoretical propositions. If we find that political information suggesting that left-wing governments will either come to or remain in power causes bond markets to switch into and/or persist longer in equilibria (regimes) connoting upward shifting yield curves in all our countries, we will have support for Proposition 2. However, if we find that such information causes bond markets to switch into and/or persist longer in upward shifting yield curve regimes in the U.K. but not in Germany, we will have support for Proposition 1.

Our forward rate data were produced in the following way. First, we generated interest rate data by estimating a theoretical yield curve. We used the method of Nelson and Siegel (1987). The resulting data are yields or spot rates for "zero coupon" Treasury securities. These yields then were

[24] For a comparison of the majoritarian and consensual features of the three countries' political systems, see Lijphart (1999) and Lijphart and Crepaz (1991). By "thick" and liquid we mean many bonds of varying maturities are regularly bought and sold in the market.

used to calculate implied forward rates. Again, the fact that the British, American, and German bond markets are thick and liquid makes these calculations meaningful. (For a discussion of the Nelson and Siegel method and the calculation of implied forward interest rates, see our Appendix.)

As for our political data, the probability of Democratic election victories was calculated using Alesina, Roubini, and Cohen's electoral option model (Alesina et al., 1997: Chapter 5, Appendix A).[25] We used Bernhard and Leblang's measure of the probability of cabinet dissolution for the U.K. and Germany. Bernhard and Leblang's (1998) probabilities are calculated from a theoretically grounded, discrete time hazard model. Note that in our period of analysis the cabinet dissolution connotes the fall of a right wing government in the U.K. and possible loss of conservative control over economic policy in Germany.[26]

Our revised EIH implies that the forward rate revisions are drawn from a mixture of two different distributions each occurring with probability $p_{t,i}$,

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[25] Alesina et al. (1997, Chapter 5) use an "electoral option model" to calculate the probability that the Democratic party will receive a majority of the two party vote (i.e., a plurality) at any point in time or

$$P_t^D = \Pr\left[V_{t+t}^{\ D} > 50\% | V_t^{\ D}; t\ ; m; s\ \right]$$

where $V^D_{t+\tau}$ is the percent who intend to vote for the Democratic party t + τ months before the election, μ is the sample mean of changes in this poll, and σ is the sample standard deviation in month to month changes in the poll. These probabilities can be calculated by the formula

$$P_t^D = \Phi\left(\frac{V_t^D + mt - 50}{s\sqrt{t}}\right)$$

where Φ is the cumulative standard normal distribution. Following Alesina et al. (1997), we use the "electoral option model" to calculate the probability of a Democratic victory in the j+d-1 months prior to an election.

[26] Given its plurality electoral system, the electoral option model could be used to calculate the expected probability of a Labour victory in Britain. However, the electoral option model assumes that the election date is known at all points in time, which is only true if the election cycle is fixed. Since this is not the case in Britain, we use Bernhard and Leblang's measure of the expected probability of cabinet dissolution instead.

$$\Delta_d Fk_{j-d,t} = N(m_i, h_{t,i}) \quad \text{with prob. } p_{t,i}, \quad i = 1,2. \tag{3}$$

In particular, we estimated the following regime switching model for each of our three countries:

$$\Delta_d Fk_{j-d,t} = m_i + \sqrt{h_{t,i}}\, e_t. \tag{4}$$

where e_t is a normally distributed *iid* error term with zero mean and unit variance and h_t is the ARCH term defined below. The subscript *i* indicates that the forward rate revision $\Delta_d Fk_{j-d,t}$ depends on an unobserved regime variable S_t which can take two values ($i = 1,2$) and evolves according to a first order Markov chain. We posit that in the first regime the yield curve is shifting upwards (indicating high inflation expectations and decreasing bond prices) while in the second regime it is shifting downwards (indicating low inflation expectations increasing bond prices)[27]. The Markov chain property implies that the probability that the process is in regime *i* at time *t* depends only on the regime it was in at time *t-1*.

$$p_{ij} = P(S_t = j \mid S_{t-1} = i) \qquad i, j = 1,2$$

Although the regimes are not directly observable it is possible to draw some inferences about them by calculating the unconditional probability that regime *i* occured at time *t* given the information set up to time *t* (Ω_{t-1}). This probability is denoted as $p_{t,i} = P(S_t = i \mid \Omega_{t-1})$.

Initally, the conditional volatility is assumed to be regime specific. In each regime *i* the conditional variance $h_{t,i}$ evolves according to an ARCH model,[28]

$$h_{t,i} = a_{i,0} + \sum_{j=1}^{p} a_{i,j} \cdot u_{t-p}^2, \tag{5}$$

[27] Once more, because a bond entitles the owner to a fixed income payment (or stream of payments) there is a negative relationship between a bond's price and its yield. Hence, when the yield curve is shifting upward, bond prices are decreasing.

where $$u_{t-1} = \Delta_d Fk_{j-d,t-1} - E_{t-2}[\Delta_d Fk_{j-d,t-1}]$$

We test whether the ARCH component of each regime is the same. The following functional form determines the transition probabilities,

$$p_{ii} = \frac{\exp(\boldsymbol{b}_{i0} + \boldsymbol{b}_{i1} \cdot x_{t-1})}{1 + \exp(\boldsymbol{b}_{i0} + \boldsymbol{b}_{i1} \cdot x_{t-1})} \qquad i = 1,2\,, \tag{6}$$

where x_{t-1} represents political information about the electoral prospects of left parties or the possibility of cabinet reorganization. Note that the functional form in (6) guarantees that the probabilities are between 0 and 1. If $\boldsymbol{b}_{i1} = 0$, then the transition probabilities are constant across time and political information has no impact on bond market equilibration. Again, our new conception of the EIH is that information about the durability of political equilibria causes switches between the different bond market equilibria represented in equations (3) – (5). (Cf. Figure 3). If this conception is accurate, the transition probabilities will depend on political information x_{t-1}. Hence we will find that the $\boldsymbol{b}_{i1}$ are statistically significant. Propositions 1 and 2 represent different theoretical expectations about the mitigating effects of institutions on the statistical significance of each country's $\boldsymbol{b}_{i1}$.

The model in equations (3) – (6) was estimated by the method of maximum likelihood as shown in Hamilton (1994). The parameters to be estimated are $m_1, m_2, \boldsymbol{a}_{10}, \ldots \boldsymbol{a}_{1p},\ \boldsymbol{a}_{20}, \ldots \boldsymbol{a}_{2p}$, $\boldsymbol{b}_{10}, \boldsymbol{b}_{11}, \boldsymbol{b}_{20}, \boldsymbol{b}_{21}$.[29]

Our actual country models were constructed in two steps. First, constant transition probability models were estimated ($\boldsymbol{b}_{11}, \boldsymbol{b}_{21} = 0$). And Wald tests were used to determine the number of regimes

[28] See Gray (1996).

[29] The likelihood function was maximized numerically by using the GAUSS constrained maximization routine.

and whether the switching process was of the simple or Markov type.[30] The number of regimes was determined by independent tests for equal means and variances. We then used Ljung-Box tests to choose the volatility specifications—parameterizations of (5)—that eliminated the serial correlation in the squared standardized residuals. Once these preliminary tests were complete, we estimated the time-varying transition probability models. It is these models that allow us to ascertain whether the transition probabilities between bond market equilibria depend on political information.

Results. We begin with the reanalysis of the American case. First, we reestimated Cohen's model using the new (Nelson-Siegel) data, the three-month revision in the forward rate on a twelve-month bond delivered nine months hence ($\Delta_3 F12_{9,t}$), and a more current sample period.[31] Once more the single equation (regime) linear regression model does not perform well. The coefficient on the variable for 3 month changes in the expected probability of electing a Democratic President, $\Delta_3 P^D_t$, is statistically significant but it has the wrong sign, implying that increases in the probability of a Democratic presidential victory causes a downward shift in the yield curve! (See Table 3, Column 1.) Not surprisingly the residual diagnostics indicate the model is misspecified. For the new time frame, Cohen's single regime linear regression model removes some but not all of the excess kurtosis in the unconditional distribution of the dependent variable. In addition, the residuals are not normally distributed and they display autoregressive conditional heteroscedasticity (Table 4).

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[30] Simple switching models are those in which the probability of being in a particular state is, unlike with the Markov switching set-up, the same regardless of the previous state.

[31] For more details refer to the appendix. The results presented in this section are robust across a number of different maturities and investment horizons. The choice of one-month vs. three-month revisions is essentially a choice about the level of temporal aggregation. Three-month revisions are quarterly revisions. Previous applications of the Markov switching model to bond markets have used quarterly data (e.g., Hamilton 1988). Because monthly series of quarterly revisions have moderately high levels of autocorrelation (e.g., see Cohen 1993, 31, Table 4.1), we report quasi-maximum likelihood standard errors in this section.

Overall, the results for our U.S. regime-switching model are much better. Our Wald tests clearly show the existence of two regimes (Table 3, Column 2). There is a downward shifting yield curve regime (regime one, $\mu_{10} < 0$) and an upward shifting yield curve regime (regime two, $\mu_{20} > 0$). The Wald test for simple switching is rejected so we have a Markov switching process where the probability of being in one bond market regime depends on the previous state of the market (fn. 30). The Ljung-Box tests show that we need an ARCH specification in both regimes to eliminate the serial dependence in our squared standardized residuals (Table 4, Column 3). Tests indicated the same ARCH model describes the conditional variance in both regimes. Note that in contrast to the single regime model, the standardized residuals from the U.S. regime-switching model do not display excess kurtosis and we cannot reject the hypothesis that the standardized residuals are normally distributed (Table 4, Column 3). Finally, we test whether both the level and three period revisions in the expected probability of a Democratic election victory affect the transition probabilities between regimes (Table 3, cols. 4,5). A Wald (t) test shows that the expected probability of a Democratic election victory has a statistically significant impact on switching from the downward shifting to the upward shifting yield curve regime. As expected, when the probability of a Democratic election victory is high, the bond market is more likely to switch out of the downward shifting yield curve regime (regime one) to the upward shifting yield curve regime (regime 2). That is, when the probability of Democratic election victory is high, the market is more likely to switch out of the regime in which traders expect low inflation into that in which traders expect high inflation. The likelihood ratio statistic for this time varying probability model is also statistically significant.[32] Figure 5 depicts the impacts of political information in the American

[32] Engel and Hamilton (1990, fn. 6) cite Gallant (1987, 219) in arguing that, for nonlinear models of this kind, likelihood ratio tests are apt to be more robust than Wald [t] tests because asymmetries in the likelihood surface can create problems for the latter type

case. The impacts are expressed in terms of basis points (fn. 10). The conditional means are those associated with the two bond market equilibria (regimes). The upper line represents the increasing yield curve associated with the condition of inflationary expectations (fn. 27); the bottom line represents the decreasing yield curve connoting the condition of deflationary expectations. The curve in the middle of Figure 5 is the unconditional mean—the expected forward rate revision, which is an average of the two conditional means weighted by the regime probabilities. Note that a large increase in the probability of a Democratic election victory (+.30) is enough to push the unconditional mean of the forward rate revision near to that of the increasing (inflationary expectation) yield curve. This is striking evidence of the impact of political information in the American case.

For the United Kingdom, our Wald tests suggest the data are again generated by two separate regimes (See Table 5). There is a downward shifting yield curve regime (regime one, $\mu_1<0$) and an upward shifting yield curve regime (regime two, $\mu_2>0$). Again, the test for simple switching is rejected. The Ljung-Box tests show that the ARCH component is enough to remove the serial correlation from the squared standardized residuals of the UK model (Table 7, Column 1).[33] Most important, we find that the probability of switching between these two British bond market regimes is significantly affected by the expected probability of cabinet dissolution. An Increase in the likelihood of cabinet dissolution raises the probability of staying in the upward shifting yield curve regime. In other words, when the expected probability that a (Conservative) British cabinet will dissolve increases, the market is more likely to remain the regime in which bond traders expect high inflation. This result is very supportive of

of test.

[33] The results in Table 5 assume the same ARCH model applies in both regimes. The same qualitative results are obtained with alternative specifications.

the EIH. The impact of political information about the possible dissolution of the Conservative cabinet on the British forward rate revision is illustrated in Figure 6. The horizontal lines again represent the mean forward rate revision for the condition of increasing (inflationary expectations) and decreasing (deflationary expectations) yield curves respectively (Table 5, cols. 3). And, the curve in the center of Figure 6 represents the impact of information about cabinet dissolution on the unconditional mean of the British forward rate revision. Note that in contrast to the American case, the impact of political information is more gradual in the British case. An increase from zero to .50 in the probability of cabinet dissolution increases the unconditional mean of the forward rate revision by about 35 basis points.

We find evidence of two bond market regimes for Germany (See Table 6). Again, there is a downward and upward shifting yield curve regime. We reject the hypothesis of simple switching; German bond market equilibration is of the Markov type. The Ljung-Box tests show that allowing for ARCH removes the serial correlation from the squared standardized residuals (Table 7, Column 2). However, in this case, none of our political variables have a statistically significant impact on the transition probabilities governing switching between the German bond market regimes. Neither the level nor change in the probability of cabinet dissolution in Germany affects the probability of a shift between increasing and decreasing yield curve regimes.

Discussion. The EIH is important in what it says about the way economic agents anticipate and act on political information. Methodologically, our analysis places the EIH on much stronger footings. The results reported here are superior to those originally produced by Alesina, Roubini and Cohen. The power of the EIH in the American case is more accurately gauged here than in that earlier work. This is because our Markov switching model has taken the excess kurtosis and ARCH in forward rate revisions into account. As our diagnostic Tables (4, 7) show, our revised EIH is sounder than Alesina, Roubini, and Cohen's EIH.

Substantively, our results provide additional confirmation of the general importance of government partisanship in relation to economic performance and of the specific impact of elections and cabinet reorganization on the workings of financial markets. Our findings are consistent with those of Garrett (1998) and many others who argue that the partisan identity of government has implications for economic policy and macroeconomic outcomes. We found that in the U.K. and U.S. the possibility of Labour and Democratic governments produced inflationary expectations and increased probabilities of higher interest rates.[34] In addition, our results add to the growing body of evidence that the workings of and prices in financial markets reflect traders expectations about the outcomes of elections and cabinet reorganizations. Our findings for government bond markets thus complement recent studies such as Herron's (2000) analysis of the impact of the 1992 British election on London Interbank Offer Rate (LIBOR) futures and stock market options and Bernhard and Leblang's (1999a) findings about the impacts of campaigns and elections on currency market efficiency.

As regards the impact of formal political institutions, our findings refute Proposition 2, the idea that the EIH is universal. Rather, our results support Garrett and Lange's (1995) argument that institutions mitigate the effect of political risk and uncertainty (Proposition 1). Our findings clearly show that the power of the EIH varies depending on the institutional context in which government bonds are traded. Our results for the case of majoritarianism and a weak central bank, Britain, are consistent with

[34] For instance, Garrett (1998) emphasizes the different propensities of left and right governments to engage in expansionary fiscal policy. A recent reanalysis of his model (King, Tomz, and Wittenber 2000, 355-57) confirms that under conditions of globalization left spending is about 2% higher than right spending.

Herron's (2000) findings for this country. Although we study a different financial asset in the context of a different model of market equilibration, we too find that Britain's institutions d not insulate financial markets from the effects of political risk and uncertainty.[35] Our results are also consistent with Bernhard and Leblang's (1998b) finding that majoritarianism provides different incentives than consensual democracy with regard to the choice of fixed versus floating exchange rates; it is politicians' appreciation for the policy flexibility afforded by majoritarianism—the very thing to which traders in British government bonds react—that makes floating exchange rates appealing in such systems.[36] Recent findings about the mitigating effects of proportional representation electoral rules vis-à-vis the workings of foreign exchange markets (Freeman, Hays, and Stix 2000) are consistent with our findings for Germany. Just as there was no evidence that the likelihood of cabinet dissolution affected the Swedish kroner/Deutsch mark exchange rate, we find no impact of this political variable on forward rate revisions of German government bonds. So it seems that in the period 1980-1995, the German's electoral system produced a strong cabinet equilibrium that reassured bond traders about the content and course of monetary policy (fn. 26). What remains to be determined is whether it is the stability

[35] On the basis of analyses of LIBOR futures and stock market futures and options, Herron (2000) estimates that if Labour had won the 1992 election instead of the Conservatives, short-term interest rates would have been about 1% higher and the British stock market would have exhibited higher volatility. Space does not permit an adequate comparison of our study and Herron's. Suffice it to say that theoretically our approaches are similar in their emphases on (rational) expectations mechanisms and substantively the magnitudes of our estimated effects of Labour rule are much alike. Methodologically however, our studies are different in that for the U.K. rather than electoral odds data we use a measure of the likelihood of cabinet dissolution (fn. 26) and we recognize the nonlinearieties and ARCH in interest rates. Herron posits normality and first-order serial correlation (2000, 329, 333). Since he does not report the relevant diagnostics in his Tables it is difficult to know if his analyses are plagued by the same problems that motivate our use of the Markov switching model here. Work on the British stock markets, for example Sola and Timmerman (1994), suggests that this might be the case.

[36] At the heart of Bernhard and Leblang's (1998b) analysis is the idea that majoritarianism allows elections and cabinet reorganizations to produce significant policy changes in a way that is not true under consensual democracy.

properties of the coalition which ruled Germany in the period of our analysis (Laver and Shepsle 1996) or larger features of Germany's consensual constitution that are of greater importance insofar as the formation of currency and bond traders' expectations are concerned. Perhaps the fortunes of

-Green coalition will give us an opportunity to sort out the effects of proportional representation in relation to other features of that country's consensual democracy.[37]

In sum, the EIH has been placed on sounder statistical footings and shown to have empirical powering certain institutional contexts. In this sense, we have refined and advanced in a significant way our understanding of how political and bond market equilibration are related.

Conclusion

The most obvious extension of our research is the incorporation of stage II or socio-economic institutions. For instance, in recent years, Iversen (1998, 1999), Hall and Franzese (1998) and others have argued that macroeconomic outcomes including inflation depend on the relations between central banks and labor market organizations. Integrating this work into our investigation presents some challenges. First, Iversen's emphasis on three kinds of wage bargaining enlarges the number of theoretically relevant cases to (3x4=) twelve. Second, the bond markets that are most critical for his argument—for example, those of Sweden and Austria—are thin and illiquid; this makes it difficult to obtain meaningful estimates of the respective forward rates. However, these bond markets are developing rapidly. In several years we should have adequate data to test some aspects of Iversen's

[37] Laver and Shepsle (1996) show that by virtue of the location of its policy preferences the conservative CDU-CSU party was likely to remain in control of economic policy regardless of major changes in the German political scene; the ruling coalition was remarkably stable in the face of various political shocks (fn. 20). To unravel the effects of consensual constitutional factors, we need cases where the ruling coalition was not stable to this extent but there are multiple veto points and other features of consensual democracy.

arguments in terms of our revised EIH. Finally, the relation between Stage II and III institutions is not well understood. Consider the Hall and Franzese article. On the one hand, our results are consistent with those of Hall and Franzese insofar as they suggest that Germany's strong central bank and "relatively centralized" labor market institutions insulate that country's bond markets from the vagaries of elector politics and cabinet equilibration. Yet, as noted above, what appears to be insulation from cabinet equlibration actually could be just the opposite; it could be that the stable cabinet equilibrium produced by Germany's proportional representation system that explains the results in Table 6. Also, Hall and Franzese argue that central bank strength alone accounts for much of the cross-national differences in inflation. But we found quite different effects of political information on bond market equilibration in the U.S. and Germany, two countries with relatively strong monetary authorities. Because formal political institutions are not yet fully incorporated in the Hall and Franzese (or Iversen) theses our results are somewhat difficult to interpret. It is these challenges for theory, design, and interpretation that will be the subject of a future paper on our new EIH.

Another extension incorporates the causal connections between countries' bond markets, that is, studying how information about the political equilibrium in one country or even in supranational government affects the equilibration of regional bond markets. In addition to the need for better data sets for small country markets, to extend the EIH in this way requires application of the technically challenging switching vector autoregressive model (Ang and Bekaert 1998).

The larger value of this paper lies in the connection it makes between political and economic equilibration in financially open democracies. It shows that in some democracies partisan politics has important macroeconomic consequences. More generally, the paper illuminates the stylized facts that a model of such democracies must explain; it charts the different ways political information about the election and cabinet prospects of left parties affect bond market behavior in different political systems. Any model of financially open democracies (e.g., cf. Freeman and House 1998) must explain why such information causes regime switching in majoritarian systems with weak central banks but not in consensual systems with strong central banks.

Appendix: Term Structure Estimation and Data Construction

To calculate the forward rate revision data used in the text we need to have expressions for forward rates or equivalently, as will be shown, for spot interest rates. Theoretically, spot interest rates could be observed from zero coupon bonds traded on various bond markets. However, since the majority of bonds are coupon bonds we need to use estimation methods to derive the spot rates empirically.[38] In this appendix we briefly review some concepts, discuss the Nelson-Siegel procedure upon which our data are based, and define the bond data used in our sample.

Concepts and Notation

The *yield curve* is a graphical depiction of the relationship between the yield on risk-free government bonds for different maturities. The *term structure* is a yield curve for zero coupon bonds; the respective yields are called *spot rates*.

In the following, let P denote the price of a coupon-bearing bond (observed market price plus accrued interest)[39], C the constant coupon payment the investor receives where the last coupon payment also includes the redemption value. The total stream of cash flows consists of $j = 1, \dots, n$ coupon payments each at the end of the j-th period. So, the bond generates a cash flow of $C_1, C_2, \dots C_n$ at maturities $m_1, m_2, \dots m_n$. Notice, that we use the following timing convention: the end of period j always corresponds to datum m_j, the end of period $j+k$ corresponds to datum m_{j+k} [40]. For simplicity, assume that the first coupon payment occurs in one period from now.

Usually, the rate of return for a j period investment is based on discrete compounding. In this context, periods usually refer to days, weeks, months, etc. If the length of the time intervals goes to zero (periods are just instants) then interest rates are compounded continuously. Thus, if the interval for which the return is measured is small, the discrete rate of return can be approximated well by the continuously compounded or log rate of return.

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[38] We gratefully thank Michael Boss for providing the data.

[39] The observed market price is called the *clean price*; the clean price plus accrued interest is called the *dirty price*.

[40] This notation states that the time interval between these two data is of length $m_{j+k} - m_j$ or k periods.

The Price of a Bond

Economic theory predicts that in a complete market the perfect foresight price of a bond is determined by the sum of the discounted future cash flows. In terms of continuous compounding, this relationship can be written as,

$$P = \sum_{j=1}^{n} C_j \cdot e^{-m_j \cdot S_j} . \tag{A1}$$

The sequence of $S_j = S(m_j)$ defines the term structure of interest rates or the zero coupon yield curve. It is the interest rate for the corresponding time-to-payment m_j.

Discount Factors

The price of a zero coupon bond is called the discount factor. Discount factors play an important role in estimating the term structure because a coupon bearing bond can be split into a bundle of zero coupon bonds, where each is discounted to the present (time t) value by means of discount functions: given a functional form for the spot rate S_j the continuous discount function which maps maturity m_j to a discount factor is defined as,

$$\boldsymbol{d}(m_j) = e^{-m_j \cdot S_j} .$$

Applying this to equation (A1) and assuming that the bond is a zero coupon bond with a redemption value of one gives

$$P = \boldsymbol{d}(m_j) , \tag{A2}$$

which shows that the discount factor for a zero coupon bond is the price of that bond.

Implicit Forward Rate

Whereas the spot rate gives the time t rate of interest for an investment until the end of period j, the *forward rate*, in general, gives the rate of interest for an k period investment undertaken some j periods in the future. Let this rate be denoted as $F_{j,k}$.

The term structure of interest rates implicitly contains information about future interest rates due to arbitrage reasons. Suppose that an investor with an investment horizon of two years faces two alternatives. She can either invest in a zero coupon bond with a maturity of two years or in a one-year bond and reinvest the proceeds in another one-year bond she will buy one year from now. Under the usual assumptions of efficient markets and rational expectations, the two alternatives should give the same overall return of investment otherwise there would be arbitrage opportunities. Formalizing this argument to arbitrary maturities it must be the case that,

$$e^{(m_j+\Delta m_j)\cdot S_{j+\Delta j}} = e^{m_j \cdot S_j} e^{\Delta m_j \cdot F_{j,\Delta j}} .$$

Rearranging in terms of $F_{j,\Delta j}$ and letting the periods become infinitesimally small ($\Delta m_j \to 0$), yields an expression for the so called *instantaneous forward rate* F_j as a function of the spot rate,

$$F_j = F(m_j) = \lim_{\Delta m_j \to 0} F(m_j, \Delta m_j) = \frac{\partial}{\partial m_j} m_j \cdot S_j . \qquad \text{(A4)}$$

In turn, solving (A4) for the continuously compounded spot rate S_j gives,

$$S_j = \frac{1}{m_j} \int_0^{m_j} F(w)dw,$$

which is the mean of the instantaneous forward rate over the interval $\left[0, m_j\right]$.[41]

Intuitively, the instantaneous forward rate gives the rate of return from buying a bond at some maturity m_j and selling it instantly. A nice interpretation of the relationship between the instantaneous forward rate and the spot rate is given in Campbell, Lo, McKinlay 1997: From the viewpoint of a borrower, the instantaneous forward rate gives the marginal cost of extending the holding period by an infinitesimal small time interval while the spot rate gives the average cost of borrowing. The relation between the two curves is therefore similar to the relation between average and marginal cost curves.

The forward rate applicable for an investment from period j to period $j+k$ is defined as the mean of the instantaneous forward rate over this interval,

$$F_{j,k} = \frac{1}{m_{j+k} - m_j} \int_{m_j}^{m_{j+k}} F(w)dw. \qquad \text{(A5)}$$

[41] Hence, the instantaneous forward rate and the spot rate are directly related. In the discussion about the discount factors, it was shown that the dicount factor and the spot rate are also directly related. Thus, knowing either one of the three measures implies the other two.

Yield to Maturity

The yield to maturity (redemption yield) R for bond i maturing in n periods is defined as,

$$P = \sum_{j=1}^{n} C \cdot e^{-m_j \cdot R} \ .$$

In contrast to the spot rates $S(m)$ which give a distinct rate for each maturity m, the yield to maturity is a summary measure, a single rate of return that produces the same present value of a bond as in the pricing equation (A1). Therefore the concept of yield to maturity is equivalent to a flat term structure where the spot rates are equal across periods and where each receipt from coupon payments can be reinvested at the same rate (the yield to maturity) over the whole investment horizon. So, the yield to maturity does not take into account that bondholders may demand different discount factors for different periods.

Notice, that the yield to maturity does not determine the price of a bond. The causality goes the other way around: the supply and demand for capital determines market clearing interest rates (the spot rates) which in turn determine the market price of a bond. Given the price, one can then calculate the yield to maturity.

Estimation

In order to explain estimation, we need to switch to a more flexible notation. Suppose there are $i = 1, \ldots, b$ bonds traded and that each bond i is generating $j = 1, \ldots, n$ payment streams occurring at maturity $m_{i,j}$ (in j periods) plus the redemption payment occurring at maturity $m_{i,n}$ (in n periods) which is again included in the last coupon payment.

To this point, it was implicitly assumed that the spot and forward rates and correspondingly the discount factors were known. In practice, one can only observe the prices P_i, the coupon payments C_i, the redemption value M_i and the timing of the cash flows $m_{i,j}$. Therefore, to estimate the term structure one has to assume a functional form either for the instantaneous forward rate, for the spot rate, or for the discount function.[42] As noted above, assuming either one of them implies the other two.

[42] This refers only to the parametric approach. Other approaches, like the spline method, are not discussed here.

Given such a functional form, the goal is to find estimates for the parameters of this function such that a distance norm between the observations and the fitted values is minimized.

Assuming a parameterization for the discount function, $\boldsymbol{d}(m_{i,j};\vec{\boldsymbol{b}})$, the pricing equation for bond *i* is given by,

$$P_i = \sum_{j=1}^{n} \boldsymbol{d}(m_{i,j};\vec{\boldsymbol{b}}) \cdot C_i \qquad \text{(A6)}$$

where the discount function is a function of maturity $m_{i,j}$ and a parameter vector $\vec{\boldsymbol{b}}$. Given that this functional specification is correct, equation (A6) must hold exactly at the true parameter vector $\vec{\boldsymbol{b}}$. However, in estimation it is assumed that the realizations of a zero mean error term $\boldsymbol{h}_i$ disturb this relationship such that

$$\begin{aligned} P_i &= \sum_{j=1}^{n} \boldsymbol{d}(m_{i,j};\hat{\vec{\boldsymbol{b}}}) \cdot C_i + \boldsymbol{h}_i \\ &= \hat{P}_i(m_{i,j}, C_i; \hat{\vec{\boldsymbol{b}}}) + \boldsymbol{h}_i \end{aligned} \qquad i = 1, \ldots, b,$$

where the parameter vector is replaced by its estimate and $\hat{P}_i$ denotes the theoretical (fitted) price of bond *i*. The norm to be minimized can be the sum of squared residuals from this nonlinear equation. Since the yield to maturity is defined by the price one can alternatively minimize the yield rather than the price errors.[43] In this case the estimation problem can be stated as,

$$R_i = \hat{R}_i(m_{i,j}, C_i; \hat{\vec{\boldsymbol{b}}}) + \boldsymbol{e}_i, \qquad \text{(A7)}$$

where $\hat{R}_i$ refers to the theoretical (fitted) yield to maturity and $\boldsymbol{e}_i$ is another zero mean error term. The problem is to find estimates of $\hat{\vec{\boldsymbol{b}}}$ such that the sum of squared residuals in equation (A7) is minimized:

$$\min_{\vec{b}} \quad SSR(\vec{\boldsymbol{b}}) = \sum_{i=1}^{b} \boldsymbol{e}_i^{\,2}.$$

[43] It is sometime argued that this is more appropriate (Schich 1997).

Nelson and Siegel's Procedure

Our estimates are based on the procedure proposed by Nelson and Siegel (1987) who assume that the relationship between the maturity and spot rates can be described by an expression which includes the sum of exponential terms. In their model, the instantaneous forward rate as a function of maturity m is modeled as,

$$F(m;\vec{\boldsymbol{b}}) = \boldsymbol{b}_0 + \boldsymbol{b}_1 e^{-m/t} + \boldsymbol{b}_2 \frac{m}{\boldsymbol{t}} e^{-m/t} ,$$

where $\vec{\boldsymbol{b}} = (\boldsymbol{b}_0, \boldsymbol{b}_1, \boldsymbol{b}_2, \boldsymbol{t})$ is a vector of parameters to be estimated. By integrating $F(m;\vec{\boldsymbol{b}})$ over the interval [0,m] and division through m, one obtains the spot rate function,

$$S(m;\vec{\boldsymbol{b}}) = \boldsymbol{b}_0 + \boldsymbol{b}_1 \frac{1-e^{-m/t}}{m/\boldsymbol{t}} + \boldsymbol{b}_2 \left(\frac{1-e^{-m/t}}{m/\boldsymbol{t}} - e^{-m/t} \right).$$

By using the relationship between discount factors and spot rates one can derive the discount function,

$$\boldsymbol{d}(m;\vec{\boldsymbol{b}}) = e^{-m \cdot S(m;\vec{b})} .$$

This functional specification can generate spot rate and forward rate curves with a variety of shapes, including upward sloping, downward sloping, hump-shaped and inverse[44].

The parametric specification of the discount function is the basis for parameter estimation by non-linear least squares. More precisely, estimation requires the following steps:

1. Select starting values $\vec{\boldsymbol{b}}^{(0)}$.
2. Calculate the theoretical spot rates $\hat{S}(m) = \hat{S}(m;\vec{\boldsymbol{b}}^{(0)})$ by making use of Nelson and Siegel's functional form for the spot rates

$$\hat{S}(m;\vec{\boldsymbol{b}}^{(0)}) = \boldsymbol{b}_0 + \boldsymbol{b}_1 \frac{1-e^{-m/t}}{m/\boldsymbol{t}} + \boldsymbol{b}_2 \left(\frac{1-e^{-m/t}}{m/\boldsymbol{t}} - e^{-m/t} \right).$$

3. Calculate the discount factors $\boldsymbol{d}(m;\vec{\boldsymbol{b}}^{(0)}) = e^{-m \cdot \hat{S}(m)}$.

[44] However, it cannot generate two local minima, etc. An extension, which adds another exponential term, has been proposed by Svensson.

4. Given the discount factors one can calculate the theoretical prices $\hat{P}_i = P_i(\vec{b}^{(0)})$:

$$\hat{P}_i = \sum_{j=1}^{n} d(m; \vec{b}^{(0)}) \cdot C_i \ .$$

5. Given the theoretical prices from step 4, one obtains the theoretical yields $\hat{R}_i$ by solving:

$$\hat{P}_i - \sum_{j=1}^{n} C_i e^{-n \cdot \hat{R}_i} = 0 \ .$$

6. Calculate the value of the objective function:

$$SSR(\vec{b}^{(0)}) = \sum_{i=1}^{b} \left(R_i - \hat{R}_i \right)^2 \ .$$

7. Check for convergence. If the criterion is met, stop. Otherwise, update $\vec{b}^{(i)}$ and repeat steps 2 to 7.

Given the final parameter estimates $\vec{b}$ and the forward rate function $F(m;\vec{b})$ one can then calculate the instantaneous forward rate for any maturity. Subsequently, given (A5), one can also calculate the mean forward rate needed for the construction of forward rate revisions.

Forward Rate Revision

Since the concept of forward rate revision deals with the *d* period change over time of the implicit forward rate for the same bond, we need to extend the notation and add time subscripts. In general, the first subscript will denote time while the meaning of the other subscripts remains the same. Applying this, $F_{t,j-1,k}$ denotes the time *t* continuously compounded mean forward rate for a *k* period bond bought in *j-1* periods (at time $t+j-1$). Similarly, $F_{t-1,j,k}$ denotes the time *t-1* forward rate for a *k* period bond bought in *j* periods (at the same time $t+j-1$).

Suppose that at time *t-1* an investor is considering to invest in such a *k* period bond which will be issued in *j* periods. The rate of return for this investment is the implicit mean forward rate $F_{t-1,j,k}$. Suppose further, that the investor decides to postpone the decision for one period. In order to evaluate this decision, she needs to compare the time *t* with the time $t-1$ forward rate for this bond. Since as of time *t*, the bond will be issued in $j-1$ periods, the appropriate forward rate for this particular bond is $F_{t,j-1,k}$. The only difference between these two forward rates is that at time t there are *j-1* periods until investment whereas as of time *t-1* there are *j* periods. Now, the one period difference between these two forward rates is called one period forward rate revision,

$$\Delta_1 Fk_{j-1,t} = F_{t,j-1,k} - F_{t-1,j,k}.$$

So, $\Delta_1 Fk_{j-1,t}$ measures the revision of the implied forward rates for the same (hypothetical, zero coupon) bond over time. Figure A1 illustrates the construction of the forward rate revisions.

Once combined with the expectations hypothesis of the term structure, the forward rate revision measures the change in the expectations of future spot rates. To show this, assume that the expectations hypothesis holds, e.g. that,

$$F_{t,j-1,k} = E_t\left[S_{t+j-1,k}\right] + \Lambda_{t,k}\,, \tag{A8}$$

where $\Lambda_{t,k}$ denotes a term premium. Equation (A8) states that the time t forward rate for a k period bond bought in *j-1* periods is equal to the time t expectations of the spot rate on a k period bond *j-1* periods ahead. Therefore, apart from the term premium, the forward rate provides the best forecast of future spot rates given time t information. Finally, by using equation (A8), the forward rate revision can be rewritten as,

$$\begin{aligned}\Delta_1 Fk_{j-1,t} &= F_{t,j-1,k} - F_{t-1,j,k} \\ &= (E_t\left[S_{t+j-1,k}\right] - E_{t-1}\left[S_{t+j-1,k}\right]) + (\Lambda_{t,k} - \Lambda_{t-1,k})\end{aligned}$$

which decomposes $\Delta_1 Fk_{j-1,t}$ into two parts: the change in the expectations about future spot interest rates and the change in the term premium. If new information arrives at time t then forward looking market participants might revise their expectations about future spot rates which influences the time t term structure.

Data

The bond data used in this paper are downloaded from Datastream Inc. which publishes a government sector bond index for all major markets including the most liquid bonds. In estimation, only those bonds listed in the index are included in the sample which guarantees a sufficient number of bonds over the maturity range where each one is fulfilling certain criteria[45].

[45] To be precise, all bonds are included in the sample that have ever been listed in the index. Datastream selects the bonds used in the index on the basis of documented criteria like turnover, maturity, etc.

For each bond, the following information is downloaded: bond ID, bond name, issue date, redemption date, coupon size, coupon dates, redemption value, and number of redemption payments. The types of bonds included are listed in Table A1; the table also contains country specific excluded bonds. Additionally, all bonds with special features (callable bonds, fungible bonds, etc.) are excluded.

Given these observations a term structure is estimated for each month in the sample by using the Nelson and Siegel procedure. With the resulting parameter estimates, we then calculate the forward rates and the forward rate revision series.

The estimation results in the paper are based on the three month revisions for a twelve month bond purchased at time t+9,[46]

$$\begin{aligned}\Delta_3 F12_{9,t} &= F_{t,9,12} - F_{t-3,12,12} \\ &= (E_t[S_{t+9,12}] - E_{t-3}[S_{t+9,12}])\end{aligned}.$$

Thus this forward rate revision measures the quarterly change in the expected twelve month interest rate that will apply in twelve months relative to time t-3 and in nine months relative to time t. In absolute time the bond will be issued at time $t+9$. According to the Fisher equation, the expected future nominal interest rates can then be split into the change in the expected real interest rates and the change in the expected inflation rates. The latter are linked to the subjective probability of the Democrats winning the next presidential elections by the rational partisan theory presented in Alesina, Roubini and Cohen. Finally this gives equation 1 in the main text,

$$\Delta_3 F12_{9,t} = (r^e_{t+9,12|t} - r^e_{t+9,12|t-3}) + (P^D_{t+9|t} - P^D_{t+9|t-3})(\boldsymbol{p}^D - \boldsymbol{p}^R) + \Lambda_{k,t}.$$

[46] At time t-3 there 12 months until the bond will be issued. At time t there are 9 months (*j*=12, *j*-3=9) until issuance.

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Table 1. Effect of Electoral Information on the Bond Market
Dependent Variable: forward rate revision ($D_dFn_{v,t}$)

Parameter	$\Delta_3F3_{7,t}$ Cohen (1993)	$\Delta_3F3_{7,t}$ (Replication)	$\Delta_3F3_{7,t}$ (Early)	$\Delta_3F3_{7,t}$ (Late)
Constant	-0.071	-0.011	0.065	-0.380
	(-0.69)	(-0.13)	(1.08)	(-0.99)
$\Delta_3F3_{7,t-3}$	-0.146	-0.087	0.098	-0.271
	(-0.90)	(-0.49)	(1.30)	(-1.04)
$\Delta_3F3_{7,t-4}$	0.095	0.107	-0.098	0.378
	(0.66)	(0.11)	(-1.07)	(1.24)
$\Delta_3\pi_{t-1}$	0.276*	0.092	0.082	0.047
	(2.10)	(1.17)	(1.68)	(0.08)
Δ_3UE_{t-1}	-0.458*	-0.399*	-0.295*	-1.241*
	(-1.97)	(-2.88)	(-2.56)	(-2.47)
Δ_3MD_{t-1}	0.102	0.009	-0.001	0.180
	(1.03)	(.281)	(-0.04)	(0.60)
$\Delta_3P^D_t$	0.027*	0.009*	0.005*	0.018
	(2.04)	(2.23)	(1.91)	(0.62)
R^2	0.11	0.06	0.07	.12
Sample	1946:12-87:2	1948:5-87:2	1948:5-79:12	1980:1-87:2
N. Obs.	329	466	380	86

Source: Cohen (1993: 47); t-statistics in parentheses; calculated with robust standard errors.
* p-value < 0.10

Table 2. Summary Statistics for Residuals from Table 1

	$\Delta_3 F3_{7,t}$ (Replication)	$\Delta_3 F3_{7,t}$ (Early)	$\Delta_3 F3_{7,t}$ (Late)
Mean	-0.002	-0.001	-0.005
Standard Deviation	1.035	0.721	1.775
Skewness	-0.023	0.613	0.705
Kurtosis	7.351	5.861	3.577
Jarque-Bera Statistic	366.776 [0.000]	152.921 [0.000]	8.221 [0.016]
$LB^2(12)$	527.058 [0.000]	162.045 [0.000]	56.891 [0.000]

Notes: Brackets contain p-values. The Jarque-Bera statistic is used to test the null hypothesis that the residuals are normally distributed. $LB^2(12)$ is the Ljung-Box Q-statistic under the null of homoscedasticity.

Table 3. Models of US Forward Interest Rate Revision ($D_3F12_{9,t}$). 4/80-12/95

Parameter/Variable	Cohen's Model	Markov Switching Model	Switching ARCH Model	Time-Varying Probabilities ($\Delta_3 P^D_{t-1}$)	Time-Varying Probabilities (P^D_{t-1})
μ_{10} (0.137)	-.147	-.741 (.104)	-.643 (.127)	-.682 (.103)	-.658 (.143)
μ_{11} / $\Delta_3 F3_{9,t-3}$	-.033 (0.160)				
μ_{12} / $\Delta_3 F3_{9,t-4}$	-.050 (0.182)				
μ_{13} / $\Delta_3 \pi_{t-1}$	-.773 (0.246)				
μ_{14} / $\Delta_3 UE_{t-1}$	-.041 (0.088)				
μ_{15} / $\Delta_3 MD_{t-1}$	-.007 (0.003)				
μ_{16} / $\Delta_3 P^D_t$	-.277 (0.113)				
μ_{20}		.624 (.164)	.366 (.123)	.325 (.100)	.326 (.112)
α_{10}	.944	.622 (.152)	.150 (.040)	.144 (.036)	.133 (.032)
α_{11}			.425 (.107)	.482 (.107)	.489 (.113)
α_{12}			.183 (.097)	.185 (.094)	.203 (.009)
α_{20}		.626 (.121)			
β_{10}		2.545 (.384)	2.445 (.392)	3.453 (.486)	2.896 (.408)
β_{11} / $\Delta_3 P^D_{t-1}$, P^D_{t-1}				-.111 (.023)	-.184 (.091)
β_{20}		1.922 (.364)	1.656 (.364)	1.521 (.364)	1.755 (.351)
β_{21} / $\Delta_3 P^D_{t-1}$, P^D_{t-1}				.018 (.021)	.004 (.011)
Log(L):	-248.29	-249.41	-216.90	-212.87	-213.60

Wald Tests

H_0: $p^{22} = 1 - p^{11}$ 256.21 [.00]

H_0: $\mu_{10} = \mu_{20}$ 108.73 [.00]

H_0: $\alpha_{10} = \alpha_{20}$ 0.00 [.98]

Likelihood Ratio Test			
	65.01 [0.00]	8.06 [0.02]	6.60 [0.04]

Parentheses contain robust standard errors. Brackets contain p-values.

Table 4. Summary Statistics for Dependent Variable and Residuals from Table 3

	Dependent Variable ($\Delta_3 F12_{9,t}$)	Cohen's Residuals	RS ARCH Residuals
Mean	-0.25	0.00	0.07
Standard Deviation	1.08	0.93	0.98
Skewness	-0.33	0.54	-0.03
Kurtosis	5.37	3.59	2.83
Jarque-Bera Statistic	47.75 [.00]	11.75 [.00]	0.26 [.88]
$LB^2(12)$	70.21 [.00]	117.11 [.00]	15.54 [.21]

Notes: Brackets contain p-values. The Jarque-Bera statistic is used to test the null hypothesis that the residuals are normally distributed. $LB^2(12)$ is the Ljung-Box Q-statistic under the null of homoscedasticity. The statistics for the regime-switching model are calculated using the standardized residuals.

TABLE 5. Regime Switching Models of Forward Interest Rate Revision ($D_3F12_{9,t}$). United Kingdom, 4/80-12/95.

Parameter	Markov Switching Model	Switching ARCH Model	Probability of Cabinet Dissolution	Δ_3 Probability of Cabinet Dissolution
μ_1	-.677	-.601 (.086)	-.596 (.079)	-.610 (.083)
(.082)				
μ_2	.866	.777 (.127)	.804 (.112)	.790 (.118)
(.135)				
α_{10}	.581 (.106)	.258 (.052)	.253 (.050)	.261 (.059)
α_{11}		.283 (.083)	.285 (.081)	.278 (.086)
α_{20}	.360			
(.083)				
β_{10}	2.172 (.303)	1.998 (.296)	2.188 (.305)	2.120 (.303)
β_{11}			-4.247 (1.075)	-4.026 (1.186)
β_{20}	1.306 (.299)	1.329 (.288)	1.324 (.296)	1.457 (.332)
β_{21}			2.191 (3.618)	2.028 (3.153)
Log(L):	-246.48	-232.39	-229.78	-229.70
Wald Tests				
H_0: $p^{22} = 1- p^{11}$	145.26 [.00]			
H_0: $\mu_1 = \mu_2$	146.43 [.00]			
H_0: $\alpha_{10} = \alpha_{20}$	3.60 [.06]			
Likelihood Ratio Test		28.17 [.00]	5.22 [.07]	5.38 [.07]

Note: Parentheses contain quasi-maximum likelihood standard errors. Brackets contain p-values.

TABLE 6. Regime Switching Models of Forward Interest Rate Revision ($D_3F12_{9,t}$). Germany, 4/80-12/95.

Parameter	Markov Switching Model	Switching ARCH Model	Probability of Cabinet Dissolution	Δ_3 Probability of Cabinet Dissolution
μ_1	-.623	-.653 (.148)	-.650 (.132)	-.643 (.174)
(.132)				
μ_2	.439	.192 (.182)	.201 (.162)	.202 (.214)
(.170)				
α_{10}	.214 (.042)	.091 (.047)	.092 (.044)	.095 (.053)
α_{11}		.429 (.163)	.418 (.150)	.419 (.174)
α_{20}	.227 (.044)			
β_{10}	2.155 (.441)	1.833 (.644)	1.846 (.676)	1.986 (.757)
β_{11}			3.275 (2.839)	.988 (1.806)
β_{20}	1.803 (.443)	1.922 (.389)	1.279 (1.015)	1.848 (.407)
β_{21}			26.186(45.881)	-.366 (.985)
Log(L):	-164.30	-154.11	-153.64	-154.10
Wald Tests				
H_0: $p^{22} = 1- p^{11}$	188.86 [.00]			
H_0: $\mu_1 = \mu_2$	165.12 [.00]			
H_0: $\alpha_{10} = \alpha_{20}$	0.04 [.84]			
Likelihood Ratio Test		20.38 [.00]	0.94 [.63]	0.02 [.99]

Note: Parentheses contain quasi-maximum likelihood standard errors. Brackets contain p-values.

Table 7. Summary Statistics for Standardized Residuals

	United Kingdom (Table 5)	Germany (Table 6)
Mean	-0.04	0.05
Standard Deviation	0.98	0.98
Skewness	0.05	0.10
Kurtosis	3.13	3.14
Jarque-Bera Statistic	0.21 [0.90]	0.49 [0.78]
$LB^2(12)$	7.64 [0.81]	15.15 [0.23]

Notes: Brackets contain p-values. The Jarque-Bera statistic is used to test the null hypothesis that the residuals are normally distributed. $LB^2(12)$ is the Ljung-Box Q-statistic under the null of homoscedasticity.

Figure 1. Excess Kurtosis in Cohen's Model's Residuals (Table 1, Column 2)

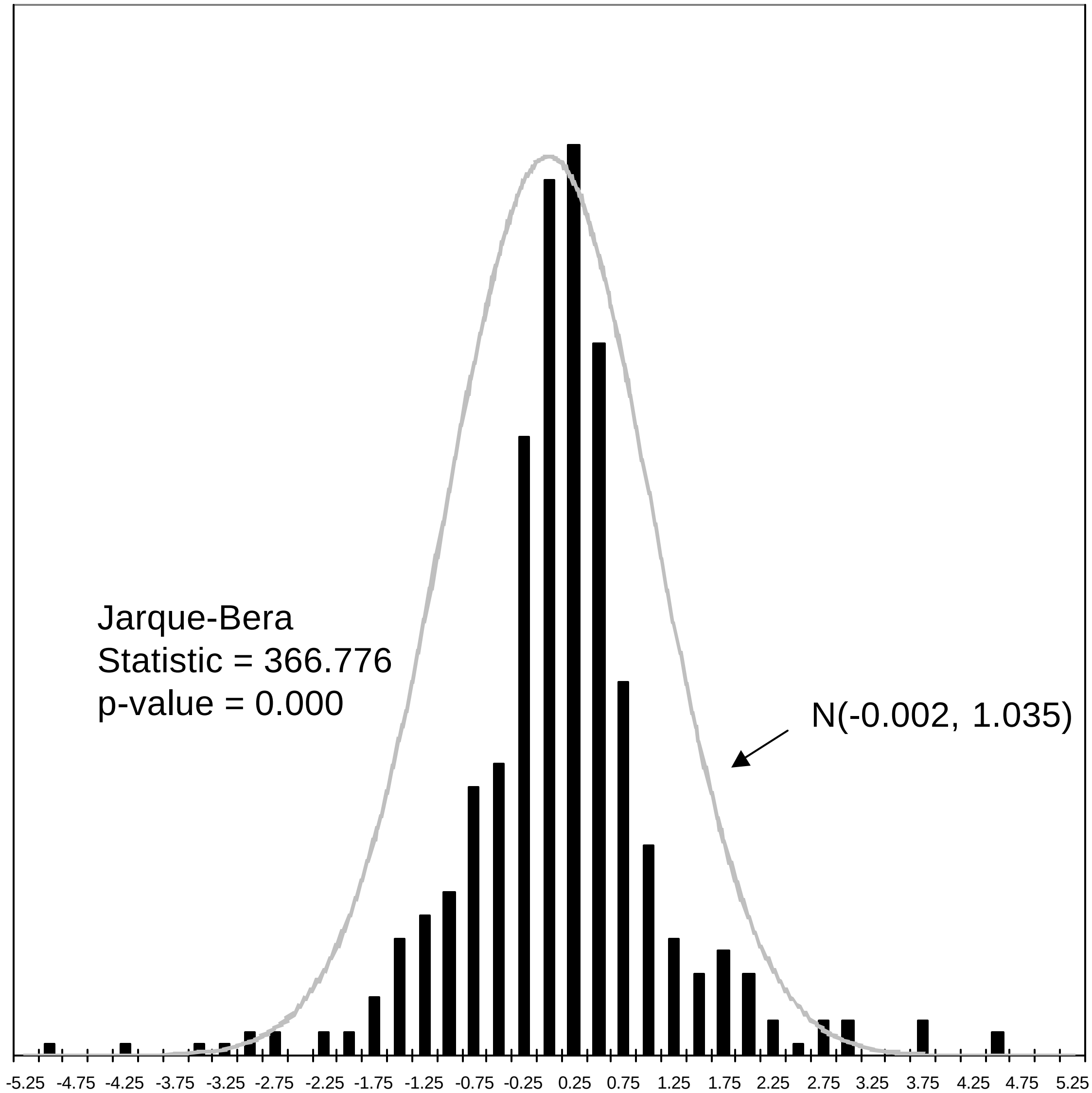


Figure 2. ARCH in Cohen's Model's Residuals (Table 1, Column 2)

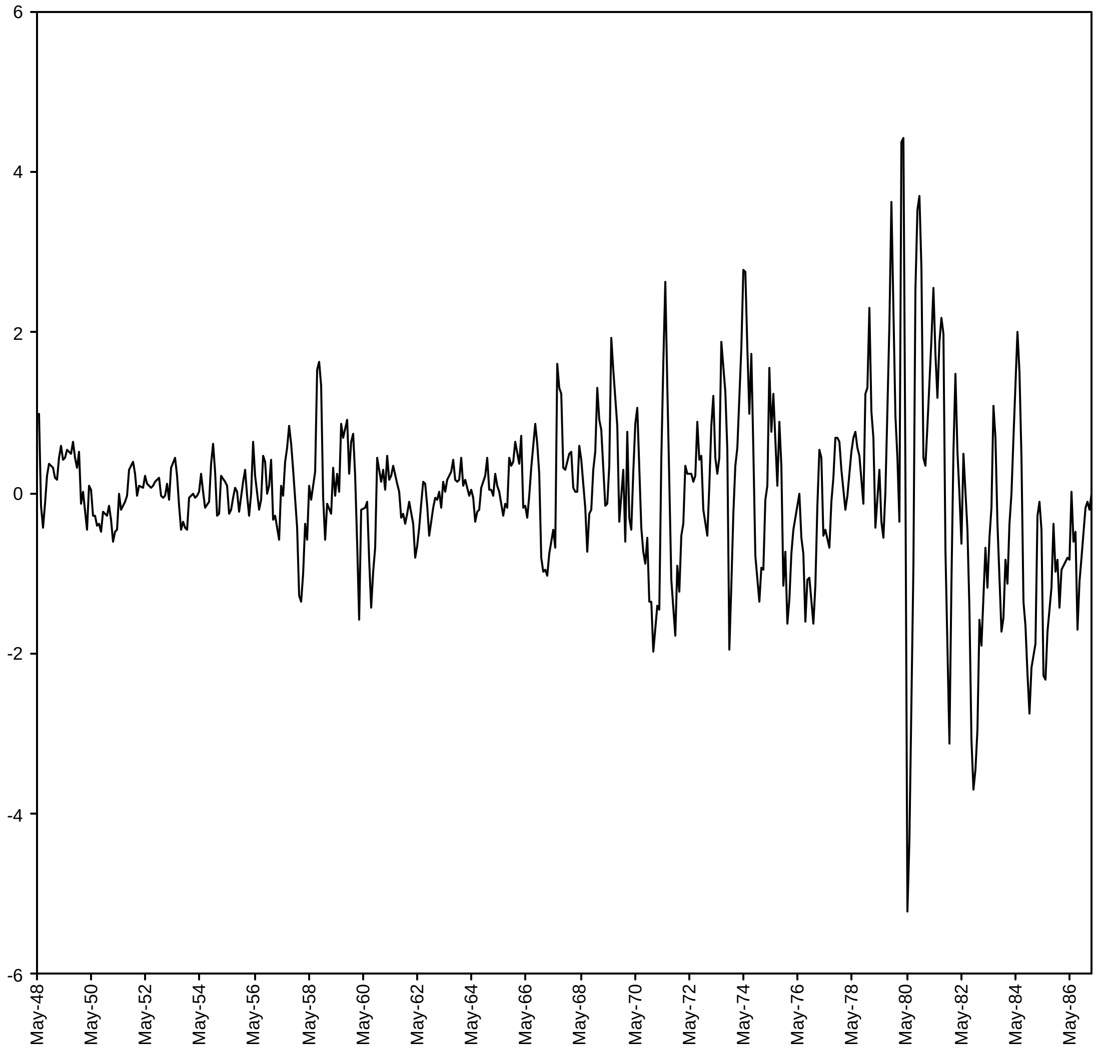
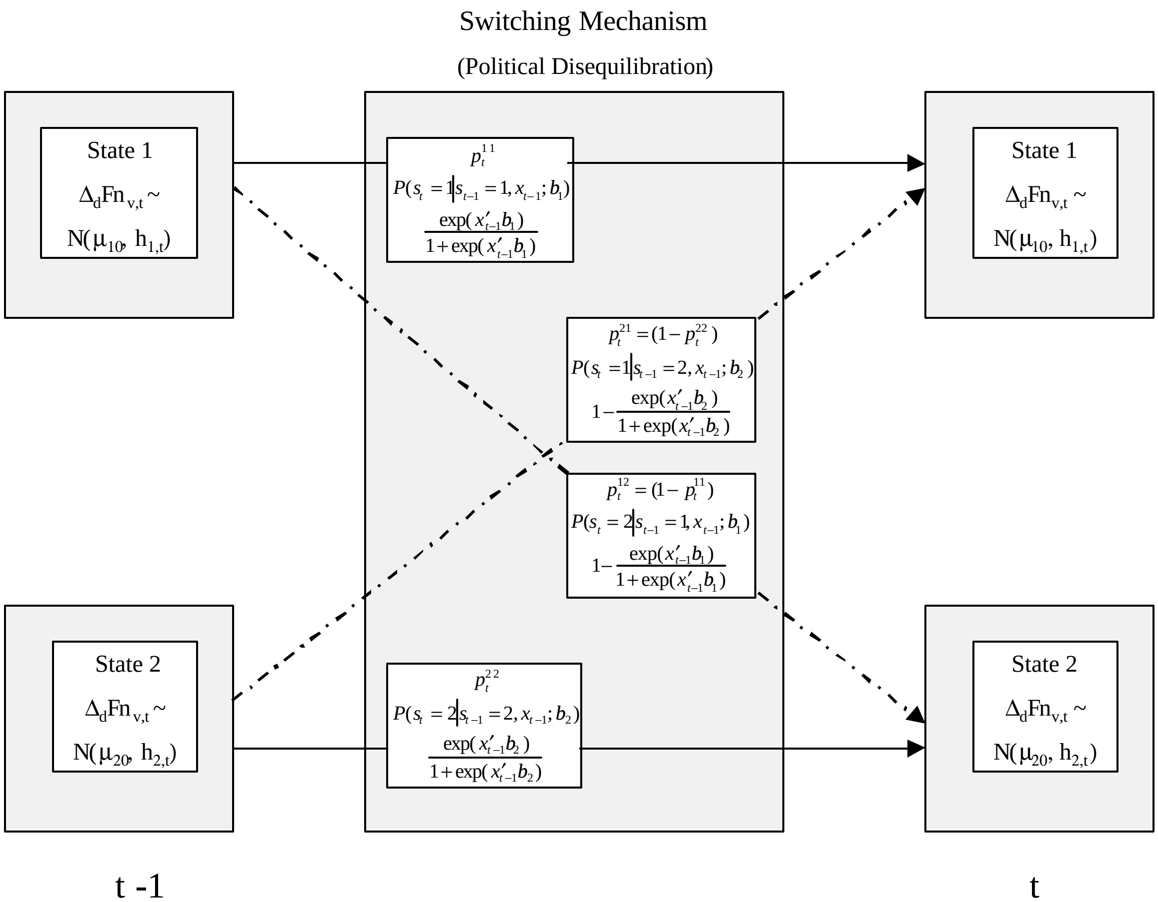
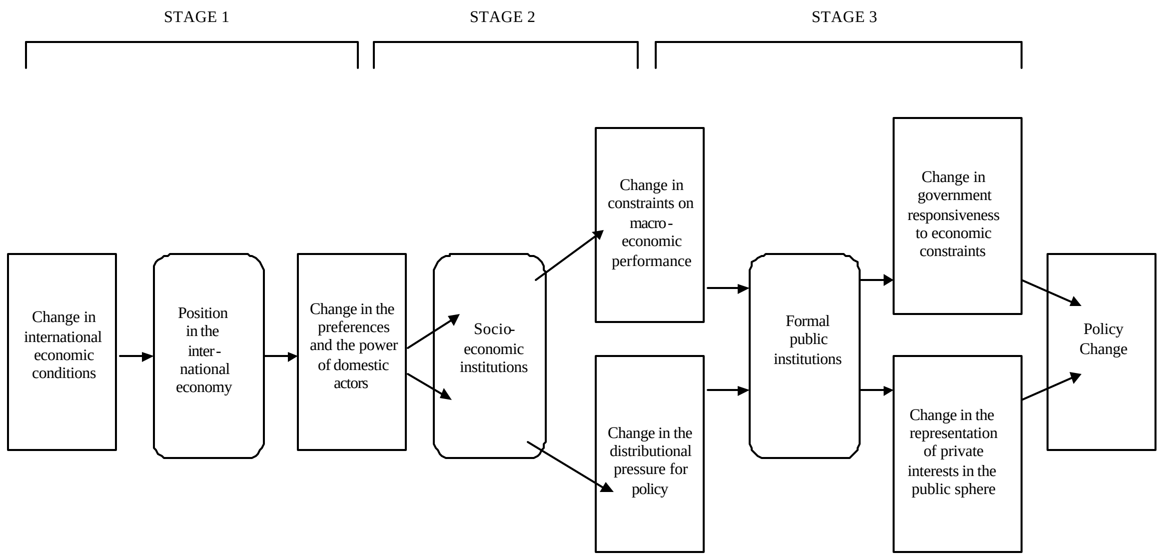


Figure 3. Markov Regime Switching Model of Forward Interest Rate Revisions



Note: $x_{t-1} = (1, x_{1,t-1}, ..., x_{(k-1),t-1})'$ and $\boldsymbol{b}_i = (\boldsymbol{b}_{i0}, \boldsymbol{b}_{i1}, ..., \boldsymbol{b}_{i(k-1)})'$, $i = 1,2$. When the last (k - 1) terms of the parameter vectors $\boldsymbol{b}_1$ and $\boldsymbol{b}_2$ are set to zero, the time varying transiton probability model collapses to the constant transition probability model. The transition probability notation is from Diebold, Lee, and Weinbach (1994, 285).

Figure 4. The International Economy, Domestic Institutions, and Political Change



Source: Garrett and Lange (1995, 630)

Figure 5. Change in the Probability of Democratic Election Victory and the Unconditional Mean of Forward Rate Revisions

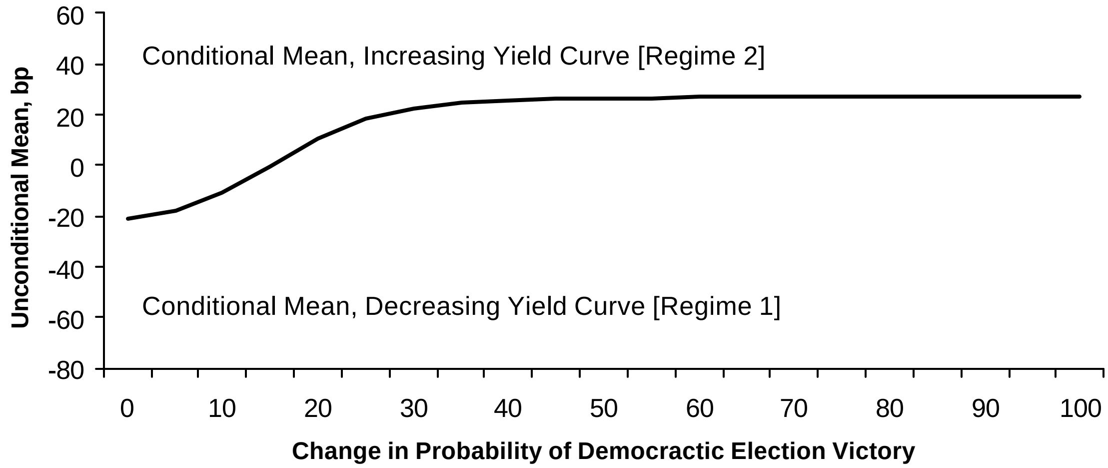


Figure 6. Probability of British Cabinet Dissolution and the Unconditional Mean of Forward Rate Revisions

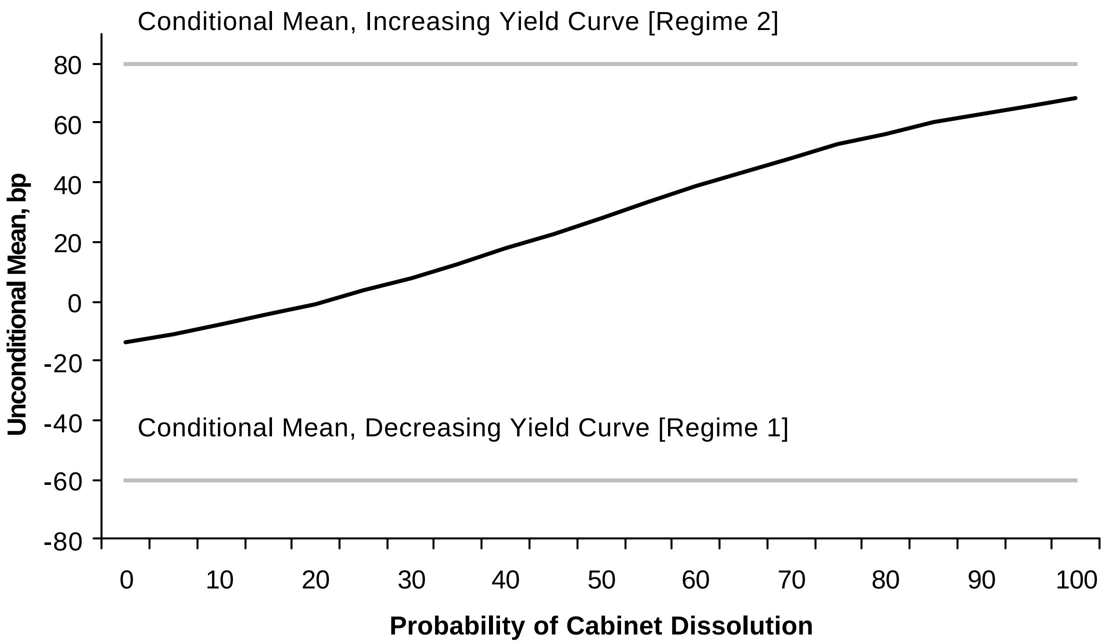
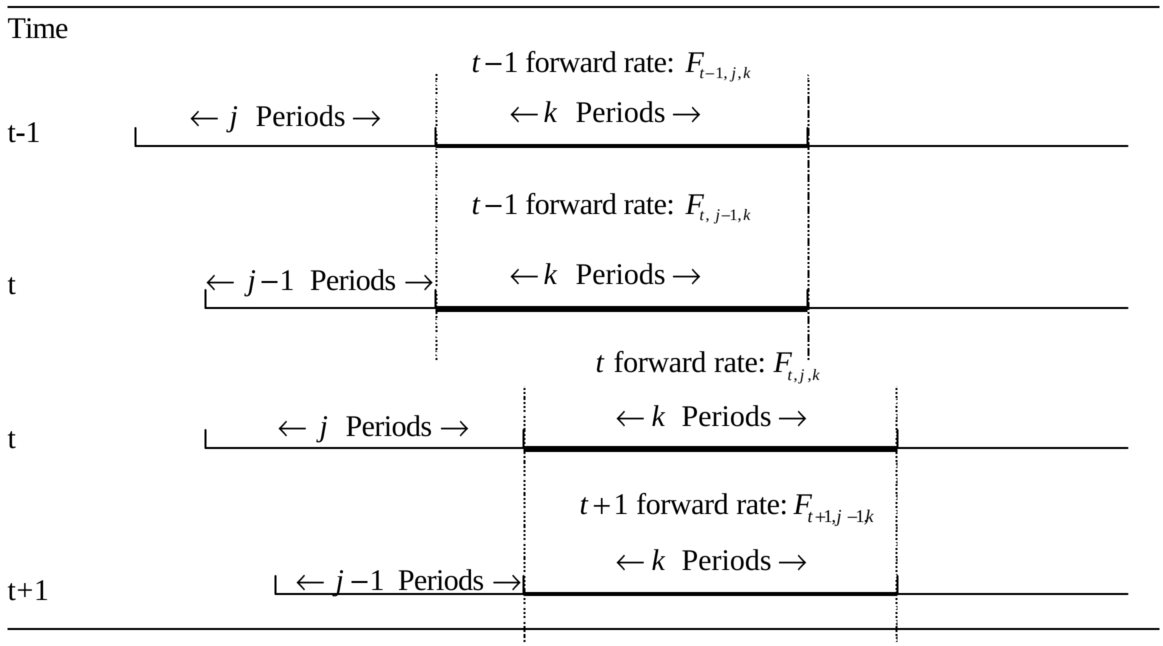


Figure A1



Time
$t-1$ forward rate: $F_{t-1,j,k}$
t-1
$\leftarrow j$ Periods $\rightarrow$
$\leftarrow k$ Periods $\rightarrow$
$t-1$ forward rate: $F_{t,j-1,k}$
t
$\leftarrow j-1$ Periods $\rightarrow$
$\leftarrow k$ Periods $\rightarrow$
t forward rate: $F_{t,j,k}$
t
$\leftarrow j$ Periods $\rightarrow$
$\leftarrow k$ Periods $\rightarrow$
$t+1$ forward rate: $F_{t+1,j-1,k}$
t+1
$\leftarrow j-1$ Periods $\rightarrow$
$\leftarrow k$ Periods $\rightarrow$