

# Coalition Government Formation and Foreign Exchange Markets: Theory and Evidence from Europe

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This study examines how the politics of coalition government formation affect foreign exchange markets in Western European parliamentary democracies. Existing studies suggest that *ex ante* uncertainty associated with the formation of coalition governments increases exchange rate volatility. We develop a formal model that places currency traders at the center and examines how traders respond to the uncertainty produced by coalition bargaining in parliamentary democracies. In sharp contrast to the literature, the model shows that traders rationally respond to uncertainty about the potential coalition government that may form and expected policies that will be implemented by the coalition government by reducing the volume of trading they do in that currency and that this, in turn, leads to a *decline* in the mean and volatility of exchange rates. Estimates from an exchange rate series of 10 Western European parliamentary democracies statistically support the claim that exchange rate volatility is negatively associated with traders' *ex ante* uncertainty about the potential coalition government that may form.

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Students of international and comparative political economy have long argued that partisanship, societal cleavages, electoral institutions, veto players, and political uncertainty stemming from elections or government formation substantially affects monetary policy, capital control policies, and the volatility of financial markets (Simmons 1994; Laver and Shepsle 1996; Clark and Hallerberg 2000; Freeman, Hays, and Stix 2000; Haggard 2000; Frieden and Stein 2001; Macintyre 2001; Bernhard and Leblang 2002a, 2002b; Clark 2002; Frieden 2002; Li and Smith 2002; Bearce 2003; Hays, Freeman, and Nesseth 2003; Kastner and Rector 2003; McGillivray 2004). As Laver and Shepsle (1996:4–5) observe:

The claim that a change of government implies at least a potential change of public policy is well illustrated by the market turbulence often caused by uncertainties about the stability of a government or the outcome of an election . . . This turbulence arises precisely because major shifts in the partisan composition of the government can imply major change in the rules of the economic game.

We explore whether uncertainty matters. Specifically, do economic actors change their behavior in response to uncertainty produced by formation of coalition

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governments in advanced parliamentary democracies and if so how does this affect foreign exchange (FOREX) markets?

Research in this issue area has essentially developed in two main directions. First, some scholars have focused on how uncertainty about electoral outcomes and uncertainty that results from change in governments in democracies affect the volatility of currency markets or commitment to certain exchange rate regimes (Freeman, Hays, and Stix 2000; Frieden and Stein 2001; Bernhard and Leblang 2002a, 2002b; Hays, Freeman, and Nesseth 2003).<sup>1</sup> Yet elections are not the only source of political uncertainty in advanced parliamentary democracies. Instead, as noted by Laver and Shepsle (1996) and Hallerberg and Von Hagen (1999), the formation of coalition governments causes substantial political uncertainty in advanced parliamentary democracies in Western Europe and this consequently affects economic policy making. In particular, Laver and Shepsle (1996:4–5, 144) emphasize that the politics associated with coalition bargaining in the process of government formation in European parliamentary democracies engenders a great deal of uncertainty about future policies among economic actors, which can lead to volatile currency markets and economic outcomes. Inspired by this literature on the “making and breaking of governments,”<sup>2</sup> the second area of research on democratic politics and financial markets studies how political uncertainty associated with government formation in advanced parliamentary democracies affect the mean and volatility of exchange rates (Lobo and Tufte 1998; Freeman, Hays, and Stix 2000; Leblang and Bernhard 2001; Bernhard and Leblang 2002a).<sup>3</sup>

Although the literature on domestic political uncertainty and currency markets has evolved in two distinct directions, it converges on a central thesis: namely, that political uncertainty associated with government formation in advanced parliamentary democracies increases exchange rate volatility and raises the risk premium in forward rates (Lobo and Tufte 1998; Freeman, Hays, and Stix 2000; Leblang and Bernhard 2001; Bernhard and Leblang 2002a). This finding has been cited as evidence against the competing idea that the inherent predictability of democratic politics, including the politics of government formation in parliamentary democracies, allows economic agents in FOREX markets to adjust and hedge against political uncertainty, which helps to preserve economic stability (Shepsle 1979; Shepsle and Weingast 1984).

While insightful, two weaknesses mar existing studies on coalition government formation and the exchange rate process. First, these studies do not provide a clear causal mechanism of how political uncertainty associated with the formation of coalition governments affects the optimal trading behavior of currency traders in FOREX markets and how this, in turn, influences the mean and volatility of exchange rates in advanced parliamentary democracies. Instead, existing studies assume that exchange rate volatility increases in advanced parliamentary democracies when uncertainty about the future coalition government that may form becomes salient because in such situations currency traders are unable to forecast the impact of government policies on the exchange rate process (Freeman, Hays, and Stix 2000:458; Bernhard and Leblang 2002a:329). The aforementioned claim is intuitive, but it fails to explain why rational economic actors in FOREX markets cannot hedge against political uncertainty by adopting trading strategies that help to minimize exchange rate volatility. Indeed, if currency traders are rational and risk averse—which is precisely what the existing literature assumes—then

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<sup>1</sup> Other studies of the impact of democratic politics on currency markets include, for example, Frieden's (2002) study of the impact of distributional preferences—between import and export-competing sectors—on variability of currency prices.

<sup>2</sup> This has been taken from the title of Laver and Shepsle (1996).

<sup>3</sup> Other scholars like McGillivray (2004) examine how uncertainty about formation of coalition governments affects stock price volatility in advanced industrial democracies.

they have incentives to adjust to political uncertainty in order to minimize volatility and dampen the effects of uncertainty on the profits they earn from currency trading.

Given the lack of a well-specified theory in the extant literature, we develop a formal model here that explicitly studies how uncertainty about the potential coalition government that may form affects the decision of currency traders in a FOREX market in advanced parliamentary democracies in Western Europe. We then examine how the response of traders to such uncertainty affects the mean and volatility of exchange rates. Our model distinguishes two types of uncertainty: *ex ante* uncertainty about (1) the potential coalition government that may emerge from coalition bargaining and (2) future policies that will be implemented by the coalition government that may form. Exploiting recent theoretical insights from the micro-structure approach to exchange rates, our model shows that when uncertainty about (1) the potential coalition government that may form and (2) future policies that will be implemented by the coalition government increases, risk-averse traders in FOREX markets have incentives to reduce the volume of domestic currency that they trade (i.e., buy and/or sell). Comparative statics from the model's solutions suggests that this leads to a decrease in the mean and volatility of the exchange rate. In contrast to the existing literature, our model thus predicts that *ex ante* uncertainty about the coalition government that may form and the policies that it might implement will *reduce* the mean and volatility of exchange rates in Western European parliamentary democracies.

The second weakness in the extant literature involves measurement of uncertainty associated with formation of coalition governments as dichotomous variables (see, e.g., Leblang and Bernhard 2001:6–7; Bernhard and Leblang 2002a) or by measuring the percentage of seats held by the coalition government (Hays, Freeman, and Nesseth 2003:215). Operationalizing political uncertainty about the potential coalition government that may form as a discontinuous dummy variable, for example, treats uncertainty about (1) government formation and (2) policy expectations as constant across all government formations. Yet in advanced parliamentary democracies there is considerable variation across government formations with respect to both the coalition that likely will form and the policies that likely will emerge. Traders use information to form expectations, and election results provide information they can use to develop their expectations about likely coalitions and the variation in policies likely to emerge from the likely coalitions. But extant research has failed to measure the differences over the different coalition bargaining situations in countries. As suggested by theories of coalition governments, traders are uncertain *ex ante* about the partisan composition and number of potential parties in the expected coalition government, the relative bargaining strength of different parties in the coalition government and the policies that will emerge from the different coalitions (Laver and Schofield 1990; Laver and Shepsle 1996; Martin and Stevenson 2001). Indicators of *ex ante* uncertainty associated with coalition government formation and future policies should measure factors such as the expected partisan composition, number of parties, and bargaining strength of different parties in the future coalition government.

In this study, we use (as described below) a more thorough and refined measure of uncertainty associated with formation of potential coalition governments and the policies that the coalition government may implement. Using these measures and data drawn from 10 Western European democracies' exchange rate to the German Deutsche Mark, we find empirically that increased uncertainty about the coalition government that could form significantly, in the statistical sense, leads to lower exchange rate volatility in 9 out of the 10 parliamentary democracies we study, but does not substantially affect the mean level of exchange rates. This finding confirms our formal model's prediction. The statistical results provide weak support for the claim that uncertainty over future policies that

will be implemented by the coalition government leads to lower exchange rate volatility.

Our study has important implications that we discuss in more detail in the conclusion, but we sketch them briefly here. First, our study reveals that as agents in currency markets automatically hedge against political uncertainty about government formation—which helps to reduce exchange rate volatility—politicians in Western European democracies ironically have low incentives to stabilize FOREX markets by reducing uncertainty associated with formation of coalition governments. This may have critical political implications that are discussed later. Second, the results presented here indicate that FOREX markets across Western European democracies are sensitive to political uncertainty that stems from the formation of coalition governments. This is important considering that prominent economists have argued that political uncertainty and politics in general do not affect exchange rates.<sup>4</sup> Third, unlike extant studies, that claim that uncertainty about government formation has adverse welfare consequences in currency markets, our study indicates that it is difficult to gauge the welfare consequences of government formation in parliamentary democracies on currency markets.

We begin the next section by presenting the formal model and the hypotheses that are derived from this model. The following section describes the generalized autoregressive conditional heteroskedasticity (GARCH) statistical model that we employ, our sample of cases, operational indicators, and the results. We conclude with a discussion of the implications of our study.

### The Model

In this section, we present a model to analyze how political uncertainty associated with coalition government formation and future policies affects the optimal amount—that is volume—of that country's currency that is bought or sold by traders. We then examine how this trading volume, in turn, influences the mean and volatility of the exchange rate process in advanced parliamentary democracies in Western Europe. We focus on the impact of political uncertainty on the exchange rate process via trading volume because a recent study by the Bank of International Settlements (2001) estimated that \$1.21 trillion/day were changing hands in these markets in 2001 and that an average volume of \$387 billion was being traded in the spot market daily.<sup>5</sup> In fact, given that the volume of trading in FOREX markets is so high, it is not surprising that macroeconomists increasingly argue that the volume of currency traded may play a key role in influencing the mean and volatility of currency prices.<sup>6</sup> Hence, a realistic formal model—which we attempt to construct here—will conceptualize how *ex ante* political uncertainty about the future coalition government that may form affect the amount of currency traders trade and how this consequently affects the exchange rate.

#### *Political Information, Fundamentals, and Exchange Rate Dynamics*

We follow recent formal work in this issue area and construct a model<sup>7</sup> where we examine how the optimal trading behavior of currency traders in FOREX

<sup>4</sup> Obstfeld and Stockman (1985) argue that it is difficult to ascertain the exact correlation between political variables and exchange rate dynamics, while in Taylor's (1995) comprehensive review of empirical tests of exchange rate dynamics in the field of economics, we did not find a single research paper in economics that addresses whether (and how) exchange rate volatility is affected by domestic political uncertainty (though see Lobo and Tufte 1998).

<sup>5</sup> Spot markets comprise the largest component (in terms of trading volume) in FOREX markets (Bank of International Settlements 2001). Our model's focus is on the interbank spot market.

<sup>6</sup> For theoretical studies on the relation between currency trading volume and exchange rate volatility in economics see Lyons (2001) and O'Hara (1995).

<sup>7</sup> The model is based on Lyons' (2001:173–175) approach to exchange rates.

markets—given political uncertainty about government formation—affects the mean and volatility of exchange rates at the macro-level. Following Lyons (2001: 173), the mean of the exchange rate in our model, denoted by  $\hat{e}_t$  and defined as the domestic price of one unit of foreign currency, follows the random process,

$$\hat{e}_t = \bar{e}_{t-1} + m_t + k_{f,bt} + k_{f,st} + \varepsilon_t. \quad (1)$$

Early theoretical work on the exchange rate process, that includes the flexible-price and sticky-price monetary models (Dornbusch 1976; Mussa 1984), suggests that macroeconomic fundamentals play a key role in determining the mean level of exchange rates. In equation (1),  $m_t$  indicates macroeconomic fundamentals that affect the mean of the exchange rate. Recent theoretical work, called the microstructure approach to exchange rates, concludes that the volume of currency traded in FOREX markets affect mean returns in currency markets (MacDonald and Marsh 1999; Lyons 2001). The parameters  $k_{f,bt}$  and  $k_{f,st}$  in equation (1) denote the volume of the currency that is bought and sold by traders. The microstructure approach to exchange rates shows that when the volume of currency traded increases (decreases), the mean spot exchange rate rises (declines) as well.<sup>8</sup> Hence, the level of trading volume is linearly related to the mean exchange rate level. As described more formally below, the traders' choice of the optimal amount of  $k_{f,b}$  and  $k_{f,s}$  is influenced by political uncertainty about the potential coalition government that may form.

The parameter  $\bar{e}_{t-1}$  in (1) is the publicly known mean of the exchange rate process from the previous period. We include  $\bar{e}_{t-1}$  because of the Efficient Markets Hypothesis, which implies that the mean level of exchange rates at any point of time reflects all available information and is merely a function of the price from the previous period (Meese and Rogoff 1983b). Finally, in (1),  $\varepsilon_t$  is a political shock to the exchange rate process. This is a critical parameter that we describe in more detail later.

We define the volatility of the exchange rate process as a random process:

$$Var(\hat{e}_t) \equiv Var(\bar{e}_{t-1}) + Var(m_t) + Var(k_{f,bt}) + Var(k_{f,st}) + \varepsilon_t, \quad (2)$$

where  $Var(\bar{e}_{t-1})$  indicates volatility from the previous period. Equation (2) is primarily based on two key approaches in the study of exchange rate volatility. Based on the monetary approach to exchange rate volatility,<sup>9</sup> we assume in equation (2) that the variability of macroeconomic fundamentals,  $Var(m_t)$ , has a linear effect on exchange rate volatility. Note that the *variability* of the volume of currency that traders trade ( $Var(k_{f,bt}), (k_{f,st})$ ) also shares a linear relationship with exchange rate volatility (Macdonald and Marsh 1999:82–83; Lyons 2001:146–147). We model it this way because in a dynamic context the variability in trading volume is characterized by a cyclical increase and decrease in the level of aggregate trading volume across time. Note from equation (1) that aggregate trading volume has a linear effect on the mean spot exchange rate. Because increasing (decreasing) trading volume leads to higher (lower) mean exchange rate levels, it also follows that the degree of change in the exchange rate is linearly affected by the degree of change—variability—in trading volume. Specifically, if trading volume changes rapidly across time—that is a situation where it is characterized by sharp dips and spikes—then the price level of the exchange rate will change rapidly as well, therein implying higher exchange rate volatility. Conversely, if the change in trading volume across time is low and stable, then price changes of the exchange rate will remain low, thus implying lower exchange rate volatility.

<sup>8</sup> When trading volume increases, it leads to appreciation of the domestic currency, which engenders an increase in the mean spot exchange rate level (Lyons 2001:173).

<sup>9</sup> Mussa's (1984) stochastic generalization of Dornbusch's (1976) exchange-rate overshooting model shows that exchange rate volatility is linearly influenced by volatility of  $m_t$ .

The most interesting component in equations (1) and (2) is  $\varepsilon_t$ , which denotes political shocks that affect the mean and volatility of exchange rates. Specifically,  $\varepsilon_t$  captures uncertainty among currency traders about the potential coalition government that may form as a result of coalition bargaining in Western European parliamentary democracies. Formally,  $\varepsilon_t$  follows a first-order autoregressive process  $\varepsilon_t = \phi\varepsilon_{t-1} + g_t$ , where  $\phi \in [0, 1]$  is a scalar known to all agents.  $g_t$  denotes the currency traders' expectation of the coalition government that may form in the near future. As there exists *ex ante* uncertainty about  $g_t$ , it is modeled as a normally distributed random variable where  $g_t \sim N(0, \sigma_g^2)$ .

In a situation where traders are uncertain about the coalition government that may form, they will also be uncertain *ex ante* about the policies that may be adopted and the effect of these policies on the exchange rate (Haggard 2000:51; Bernhard and Leblang 2002a). *Ex ante* uncertainty about future policies among traders may occur, for example, because of their inability to predict *ex ante* the policies that may be implemented by the future coalition government. At the same time, however, scholars and policy analysts have noted that in Western European democracies, politicians from parties that are likely to form the coalition government and occupy key ministerial positions in the coalition government often make "announcements" about monetary and fiscal policies that may be implemented in the future (Kritzinger, Klemmensen, and Chari 2004). Therefore, when coalition bargaining occurs, traders will observe a signal,  $S_t$ , comprised of policy announcements made by politicians. Currency traders will use the information acquired from this signal to predict how the one-period ahead exchange rate will evolve as a function of their expectation about the type of policies that may be adopted by the potential coalition government that may form.

As the announced policies that may be adopted will depend on the potential coalition government that may form, we define the signal as a function of  $g_t$  and a parameter  $\psi_t$ ,  $\alpha S_t = g_t + \psi_t$ . In a situation of coalition government formation, traders are uncertain *ex ante* about whether the announced policies may be implemented or not; hence the parameter  $\psi_t$  in  $S_t$  is normally distributed, that is  $\psi_t \sim N(0, \sigma_\psi^2)$  to formalize the intuition that the signal provides information about expected policy, albeit with some uncertainty.  $\alpha \in [0, 1]$  denotes the weight that traders place on acquiring information from the signal  $S_t$ . From  $\alpha S_t = g_t + \psi_t$ , the expected political shock conditional on the signal is defined as  $E(\varepsilon_t | S_t) = \phi\varepsilon_{t-1} + \alpha S_t$ . Given  $E(\varepsilon_t | S_t)$ , the one-period ahead forecast of the exchange rate, conditional on information from  $S_t$  is

$$E(\hat{e}_t | S_t) = \bar{e}_{t-1} + \phi\varepsilon_{t-1} + \alpha S_t. \quad (3)$$

As described more formally below, currency traders take the *signal extraction formula* in (3) into account before choosing the optimal volume of currency that they buy or sell.

#### Optimal Trading Behavior of Currency Traders

To keep the model simple, we study how uncertainty associated with coalition government formation in European parliamentary democracies affects the optimal trading behavior of a representative currency trader (*i*) who acts as a proxy for a continuum of identical traders. The currency trader's optimization problem is to choose an optimal amount of currency to either buy or sell such that it maximizes his (or her) profit. As a buyer, the trader faces an inverse, nonlinear demand function,  $a - \frac{(k_{f,b})^2}{2}$ , where  $a > 0$  and  $k_{f,b}$  the volume of currency that he seeks to buy. Assuming that the exchange rate is measured in the German DM, the cost of buying in terms of the traded currency will be given by  $\hat{e}k_{f,b}$ . The trader's profit function when s/he buys is thus,

$\pi_B = (a - \hat{e})k_{f,b} - \frac{(k_{f,b})^2}{2}$ . At each point of time, the trader decides how much to buy, which depends in part on the exchange rate level expected to prevail over the next period. Based on existing microstructure models of exchange rate dynamics (O'Hara 1995; Lyons 2001), the trader (as a buyer) maximizes the following expected mean-variance utility function, which is increasing in expected profits and decreasing in the variance of profits, conditional on  $S_t$ ,

$$E_i(U_B|S_t) = E(\pi_B|S_t) - \frac{1}{2}\gamma_B \text{Var}(\pi_B|S_t), i = 1, 2, \dots, n, \quad (4)$$

where  $\gamma_B$  denotes the coefficient of the Arrow-Pratt measure of absolute risk aversion  $0 < \gamma_B < \infty$  for the buyer. The full form of the expected profit function, conditional on  $S_t$  is obtained from substituting the signal extraction formula  $E(\hat{e}_t|S_t) = \bar{e} + \phi\varepsilon_{t-1} + \alpha S_t$  for the parameter  $\hat{e}$  in the profit function  $\pi_B$ . Completing the substitution, we obtain

$$E_i(\pi_B|S_t) = (a - (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t))k_{f,b} - \frac{(k_{f,b})^2}{2}. \quad (5)$$

Note that the variance component in (4) ( $\text{Var}(\pi_B|S_t)$ ) is defined as  $E(\pi_B - E(\pi_B|S_t))^2$ . Substituting equation (5) into  $E(\pi_B - E(\pi_B|S_t))^2$ , we obtain, after some algebra, the full form of the variance of the profit function conditional on  $S_t$ ,

$$\text{Var}_i(\pi_B|S_t) = E(\pi_B - E(\pi_B|S_t))^2 = (\alpha\sigma_\psi^2 + \sigma_m^2)(k_{f,b})^2. \quad (6)$$

As a seller, the currency trader incurs costs for procuring the amount of currency ( $k_{f,s}$ ) required to sell to other traders or customers (Lyons 2001). These costs could involve the transaction costs of trading or the costs of borrowing capital to procure the exact volume of currency required for purposes of selling (O'Hara 1995). Without loss of generality, we assume that the trader incurs a nonlinear cost function  $ck_{f,s} + \frac{(k_{f,s})^2}{2}$  where  $c > 0$  is the marginal cost parameter and  $k_{f,s}$  denotes the amount of currency sold. Normalizing the price of  $k_{f,s}$  in foreign currency to unity, the sellers' random profit in terms of the price of the domestic currency is  $\pi_s = (\hat{e} - c)k_{f,s} - \frac{(k_{f,s})^2}{2}$ . Given  $\pi_s$ , the seller maximizes the following expected mean-variance utility function, which is increasing in expected profits and decreasing in the variance of profits, conditional on  $S_t$ ,

$$E_i(U_s|S_t) = E(\pi_s|S_t) - \frac{1}{2}\gamma_s \text{Var}(\pi_s|S_t), \quad (7)$$

where  $\gamma_s$  denotes the coefficient of absolute risk aversion  $0 < \gamma_s < \infty$  for the seller ("s"). Substituting the signal extraction formula for  $\hat{e}$  in  $\pi_s$  yields the full form of the expected profit function, conditional on the signal  $S_t$ ,

$$E_i(\pi_s|S_t) = (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t - c)k_{f,s} - \frac{(k_{f,s})^2}{2}. \quad (8)$$

The full form of the variance component of the profit function for the seller is, therefore,

$$\text{Var}_i(\pi_s|S_t) = E(\pi_s - E(\pi_s|S_t))^2 = (\alpha\sigma_\psi^2 + \sigma_m^2)(k_{f,s})^2. \quad (9)$$

#### *Comparative Statics and Testable Hypotheses*

In the appendix, we formally characterize the optimal volume of the currency that is bought ( $k_{f,b}^*$ ) and sold ( $k_{f,s}^*$ ) as well as the variability of the volume of currency that traded, ( $\text{Var}(k_{f,b})$ ) and ( $\text{Var}(k_{f,s})$ ). Comparative statics on the aforementioned solutions provide the following substantive result,

**Proposition 1:** *When uncertainty about the potential coalition government that may form*

increases ( $\sigma_g^2$ ) in Western European parliamentary democracies, the mean level ( $\hat{e}_t$ ) and volatility of the exchange rate strictly decreases.

*Proof:* See Appendix A ■

The causal intuition that explains the result in Proposition 1 is as follows. First, when uncertainty about the potential coalition government that forms increases, the political shock to the exchange rate process—conditional on the signal  $S_t$ —becomes negative (i.e.,  $\partial E(\varepsilon_t|S_t)/\partial\sigma_g^2 < 0$ ). A negative shock leads to a decrease in the mean exchange rate because in the signal extraction formula in (3), the mean of the exchange rate process conditional on  $S_t$  is linearly related to the political shock. Second, because currency traders in the FOREX market are risk averse, they rationally respond to higher political uncertainty about the potential coalition government that could form by buying and selling (in other words, trading) a lower amount of the foreign currency ( $\partial k_{f,b}^*/\partial\sigma_g^2 < 0$ ,  $\partial k_{f,s}^*/\partial\sigma_g^2 < 0$ ). This leads to a decrease in the volume of the foreign currency that is traded. As the mean level of the exchange rate,  $\hat{e}_t$ , is linearly dependent on the volume of the traded foreign currency—see equation (1)—a decline in the volume of trading leads to a decrease in the mean exchange rate level *ceteris paribus*.

Two reasons explain why exchange rate volatility decreases when uncertainty about the coalition government that may form increases. First, risk-averse traders rationally respond to higher uncertainty about the potential coalition government that may form by reducing the variability of the volume of the traded currency ( $\partial\text{Var}(k_{f,b}^*)/\partial\sigma_g^2 < 0$ ,  $\partial\text{Var}(k_{f,s}^*)/\partial\sigma_g^2 < 0$ ). This is so for two reasons. First, as traders are risk averse, they dislike volatility associated with trading. Second, observe from the mean-variance utility functions in (4) and (7), that higher variability in the volume of trading will decrease the traders' expected utility. Therefore, as mentioned above, risk-averse traders have rational incentives to reduce variability in the volume of trading. Now, we know from equation (2) as well as existing studies that exchange rate volatility shares a direct linear relationship with the variability of the amount of currency that is bought and sold in the FOREX market. Hence, when variability of the volume of the traded currency falls, it leads to a decline in exchange rate volatility, as claimed.

We also obtain the following result from our model,

**Proposition 2:** *When uncertainty about expected policy that will be implemented by the potential coalition government that may form increases ( $\sigma_\psi^2$ ) in Western European parliamentary democracies, the mean and volatility of the exchange rate decrease.*

*Proof:* See Appendix A ■

The intuition driving the finding in Proposition 2 is as follows. Our model shows that when uncertainty about future policies that will be implemented by the potential coalition government increases, the variance of the profit earned by traders increases as well.<sup>10</sup> When the variance of the traders' profit increases, risk-averse traders again have incentives to decrease the variability of the volume of the traded currency ( $\partial\text{Var}(k_{f,b})/\partial\sigma_\psi^2 < 0$ ,  $\partial\text{Var}(k_{f,s})/\partial\sigma_\psi^2 < 0$ ) because higher variability in trading volume decreases their expected profit. Following equation (2), lower variability in the volume of the traded currency leads to lower exchange rate volatility. Further, when uncertainty over future policies increases, risk-averse traders rationally reduce their volume of trading. A drop in the volume of trading leads to a decline in the mean exchange rate level. Propositions 1 and 2 from our formal model thus provide the following testable hypotheses:

<sup>10</sup>  $\partial\text{Var}(\pi_B|S_t)/\partial\sigma_\psi^2 = \alpha^2 k_b^2 > 0$  and  $\partial\text{Var}(\pi_s|S_t)/\partial\sigma_\psi^2 = \alpha^2 k_s^2 > 0$ .

**Hypothesis 1:** *When uncertainty about the potential coalition government that may form increases, the mean and volatility of the exchange rate decreases in Western European parliamentary democracies.*

**Hypothesis 2:** *When uncertainty about the expected policy that will be implemented by the potential coalition government that may form increases, the mean and volatility of the exchange rate decreases in Western European parliamentary democracies.*

### Statistical Methodology

To test Hypotheses 1 and 2, we employ a generalized autoregressive conditional heteroscedasticity model (GARCH) with an autoregressive term in the conditional mean equation (Bollerslev 1986). The advantage of a GARCH(1, 1) model is that it is comprised of two equations: one for the conditional mean and one for the conditional variance. This is important because our theoretical model provides predictions on how uncertainty associated with coalition government formation affect the mean and volatility of the exchange rate.

The GARCH model is appropriate for our tests for three reasons. First, in our formal model, we examine the impact of political uncertainty on two continuous variables: the degree of the mean and variance of exchange rates. Unlike Freeman, Hays, and Stix (2000) and Hays, Freeman, and Nesseth (2003), we do *not* examine whether political uncertainty engenders a “switch” from a state of high volatility to low volatility or vice versa. Thus, it is more appropriate to estimate GARCH rather than Markov–Switching models.<sup>11</sup> Second, it can account for the large clustering of errors (i.e., volatility clustering), which are common in exchange rate series. Third, estimates from Engle’s (1982) Lagrange–Multiplier (LM) test indicate the presence of autoregressive conditional heteroscedasticity (ARCH) in the residuals in each of the monthly exchange rate series that we examine.<sup>12</sup> This necessitates the use of GARCH models (Engle 1982; Bollerslev 1986). The conditional mean equation for a GARCH (1, 1) model with a set of exogenous variables can be written as

$$\Delta y_t = \xi + \lambda Z_t + \varepsilon_t, \quad (10)$$

where  $\Delta y_t$  is the change in the observed spot exchange rate,  $\xi$  is a constant,  $Z_t$  is a vector of exogenous political and economic variables and the error term is  $\varepsilon_t \sim N(0, \sigma_t^2)$ . In GARCH models, we can also specify how the conditional variance evolves ( $\sigma_t^2$ ) over time in response to the value of past variance, prior shocks, and exogenous political and economic variables. The conditional variance equation for the standard GARCH (1, 1) model is

$$\sigma_t^2 = \omega + \theta_1 \varepsilon_{t-1}^2 + \beta_1 \sigma_{t-1}^2 + \Phi I_{i,t}. \quad (11)$$

$\sigma_t^2$  is the one-period ahead forecast variance based on information at time  $t - 1$ .  $\sigma_t^2$  is a function of: the constant ( $\omega$ ), the ARCH term ( $\varepsilon_{t-1}^2$ ), the GARCH term ( $\sigma_{t-1}^2$ ), and a set of exogenous political and economic variables ( $I_{i,t}$ ). We include an AR(1) term in the mean equation to correct for serial correlation and to check whether there exists a statistical correlation between the lag and the current mean exchange rate level, as predicted by the Efficient Markets hypothesis. As the Jarque–Bera test indicates that kurtosis exists in the exchange rate series we use, the GARCH model is estimated with Bollerslev and Wooldridge (1992) semi-robust standard errors that provide unbiased standard errors that are robust to deviations in normality of

<sup>11</sup>Hansen’s (1992) test for each European country’s currency in the price of German DM on a monthly basis—that we examine—fails to reject the null of *no* switching in the mean and variance at the 1% level. Hence, we do not need to estimate Markov–Switching models.

<sup>12</sup>The  $p$ -value from Engle’s LM test for each exchange rate series that we examine is approximately equal to zero, which means that we can reject the null of homoscedasticity.

the residuals. The results from the GARCH models remain robust when we use the Student  $t$  or GED distributions to reduce kurtosis.<sup>13</sup>

### Sample and Dependent Variable

We collected relevant data for a sample of 10 Western European parliamentary democracies from 1971–1989. This includes the following democracies with a PR (or semi-PR) electoral system and a history of coalition governments: Austria, Belgium, Denmark, France, Ireland, Italy, Luxembourg, Netherlands, Norway, and Sweden. We discuss below why we chose this particular sample for our empirical tests.

To measure our dependent variable—the mean and volatility of the spot exchange rate—we obtained time-series data on monthly spot exchange rates for each of the 10 democracies mentioned above. Each exchange rate is measured as the price (in German DM) of 1 unit of foreign currency. Therefore, our data consist of 10 exchange rate series: German DM/Austrian Schilling (1972–1988), DM/Belgian Franc (1971–1989), DM/Danish Krone (1972–1984), DM/French Franc (1971–1988), DM/Luxembourg Franc (1971–1987), DM/Irish Punt (1971–1989), DM/Italian Lira (1971–1988), DM/the Netherlands (Dutch) Guilder (1972–1988), DM/Norwegian Kroner (1971–1989), and DM/Swedish Krona (1971–1988).<sup>14</sup> We use the German DM as the benchmark currency as the DM was the most important anchor currency for Western European governments from 1971 to 1989.

To estimate the GARCH(1, 1) model in equations (10) and (11), we used the first difference on the monthly spot exchange rate in each sample.<sup>15</sup> Differencing the monthly spot exchange rate ensures that each exchange rate series is nonintegrated.<sup>16</sup> The data for the monthly spot exchange rate in each case are taken from the IMF's *Government Financial Statistics* CD-Rom. We also estimated the GARCH model on the first difference of each of the monthly spot exchange rates series with respect to the U.S. \$. We do not report the results with respect to the U.S. \$ owing to a lack of space, but these results are available on request.

We restrict our sample to time-series data from the 10 Western European democracies because of four reasons. First, note that our theoretical framework addresses how the politics of coalition government formation affects exchange rates in well developed and technologically advanced currency markets where trading volume is sensitive to political information in the short-run. Technologically advanced and sophisticated currency markets are common to only advanced industrial democracies (i.e., OECD countries), therein explaining why we focus on Western European parliamentary democracies. Second, observe that among advanced industrial democracies, coalition governments largely form in parliamentary democracies with a Proportional Representation (PR) or semi-PR electoral system. As most Western European parliamentary democracies are PR or semi-PR countries, where the formation of coalition governments is the norm rather than the exception (see Table 1), we thus empirically focus on the Western European democracies listed above. Focusing on the 10 countries mentioned earlier also helps us to control

<sup>13</sup> Estimates from the GARCH models with Student  $t$  or GED distributions are available on request.

<sup>14</sup> As described below, data for our political variables—taken from Martin and Stevenson's (2001) empirical study of government formation—is available till 1984 for Denmark, till 1988 for the Netherlands, Italy, and Sweden and till 1989 for the remaining countries.

<sup>15</sup> As is well known, exchange rate series are integrated (i.e., have unit roots) and are thus subject to spurious regression. Diagnostic tests on these series confirm the presence of unit roots, and thus we use the first difference of the series.

<sup>16</sup> The Phillips-Perron unit-root test rejects the existence of a unit root for each of the five differenced exchange rate series at the 1% level.

TABLE 1. Coalition Governments in Western European Democracies in Sample

	1945–1999				Average # of Parties in Government
	Coalition Governments (n)	%	Single Party (n)	(%)	
Austria	16	76.2	5	23.8	1.8
Belgium	28	84.8	5	15.2	3.2
Denmark	17	54.8	14	45.2	1.9
France	17	73.9	6	26.1	2.3
Ireland	10	45.5	12	54.5	1.6
Italy	34	69.4	14	28.6	2.9
Luxembourg	16	100	0	0	2.1
Netherlands	22	100	0	0	3.2
Norway	8	30.8	18	69.2	1.7
Sweden	7	26.9	19	73.1	1.4

*Notes:* Data for coalition governments and the average number of parties in government for the nine European democracies from 1945 to 1999 is taken from Mitchell (forthcoming) and Martin and Stevenson (2001).

for system type via design because all these countries are PR systems and parliamentary democracies.

Third, the continuous measure of uncertainty associated with government formation that we employ (described below) requires detailed data on factors such as the ideological divisions within the potential coalition, the ideological divisions within the opposition (relevant for minority coalitions only), whether the party of the previous prime minister is a member, whether an investiture rule is required (relevant for minority governments only), and whether the coalition contains a party that is “very strong” or “merely strong” (à la Laver and Shepsle 1996). Unfortunately, detailed data on the aforementioned factors are only available for the 10 countries that are being examined here. Fourth, our sample partly coincides with the data used by Leblang and Bernhard (2001) and Freeman, Hays, and Stix (2000). For example, Leblang and Bernhard (2001) examine volatility of the Belgian Franc and Swedish Krona series between 1980 and 1996, while Freeman, Hays, and Stix (2000) examine volatility of the Deutschmark series. To some extent, this allows us to compare our results more fruitfully to those obtained by Leblang and Bernhard (2001) and Freeman, Hays, and Stix (2000).

That said, there are two main limitations in our sample. First, the temporal unit of observation in our sample is *monthly* spot exchange rates. We would have preferred a more fine-grained aggregation (e.g., week or day), but our primary independent variables of interest do not vary at lower levels of aggregation. A finer frequency would also preclude use of standard macroeconomic fundamentals such as inflation and output that are unavailable at higher frequencies. In addition to monthly observations, observe that the temporal domain of our sample ends in 1989. This is because the relevant political data that have been used to construct our measure of uncertainty associated with government formation is available on a comprehensive basis only from 1971 to 1989 (or 1988) for each of the 10 European democracies that we examine here.<sup>17</sup> The data required to operationalize our continuous measure of uncertainty are only available for three countries in our sample from 1989 to 1999. Because we prefer to keep our sample as large as possible to enhance generalizability of the results, we did not extend the temporal domain of our sample to the 1990s since doing so will force us to use a much smaller sample.

<sup>17</sup> The data for our political variables are largely taken from Martin and Stevenson’s (2001) empirical study of government formation for the European democracies examined here.

Although the lack of sufficient time-series data that extends beyond 1989 will not directly threaten our inferences, we believe that the relatively short sample used here limits our empirical analysis in two main ways. First, our smaller data set may not contain substantial variation in political uncertainty associated with coalition government formation. As a result, we may actually be *underestimating* the impact of our key independent variable—uncertainty associated with coalition government formation—on exchange rate volatility. This could have been avoided with a larger data set, but as mentioned above, we simply lack detailed data on critical variables that are required to construct our continuous measure of uncertainty. The second problem is that the empirical inferences that we derive from the smaller sample are clearly not as generalizable in the cross-sectional sense, as we would ideally desire. Hence, in future research, we hope to include other Western European parliamentary democracies in our sample such as Finland and Iceland for which we currently lack the necessary data to operationalize our measure of political uncertainty. We also hope to extend the temporal domain of our sample in future research.

#### *Independent and Control Variables*

A thorough test of our formal model's causal mechanism would require us to test the interactive effect between uncertainty about government formation and trading volume on the mean and volatility of exchange rates. We had hoped to do so, but unfortunately we could not because data on the volume of currency traded for each series that we examine is unavailable for the temporal period being examined here. Institutions such as the Bank for International Settlements (BIS), the IMF, World Bank, and OECD which maintain some data on the volume of currencies traded in American and international FOREX markets do not have volume data for any of the exchange rate series analyzed here during the 1970s and 1980s.<sup>18</sup> An exhaustive search on exchange rate data held by large private institutions also revealed that volume data are simply not available for the 1970s and 1980s. Hence, following standard convention in the literature, we test the correlation between uncertainty about government formation and the exchange rate process as posited in Hypotheses 1 and 2.

To test Hypothesis 1, we need a measure of political uncertainty that stems from the dynamics of coalition bargaining during the process of government formation. Following the procedure described by Martin (2004), we construct an index of “bargaining uncertainty” that measures the dispersion of formation probabilities across potential governments in a bargaining situation.<sup>19</sup> This index is constructed in the following way:<sup>20</sup> First, we used the parameter estimates from the Martin and Stevenson's (2001) study of government formation (the estimates from their final model, Model 10) to generate predicted probabilities ( $P_{ij}$ ) of formation for every coalition in every bargaining situation. Specifically, in the conditional logit estimated by Martin and Stevenson (2001), the probability that in bargaining situation

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<sup>18</sup> The Bank For International Settlements (BIS) maintains some publicly available data on global foreign exchange market turnover—i.e., trading volume of various currencies—in different FOREX markets. A thorough search of BIS' databases, working papers and statistics revealed references to daily and monthly trading volume data for the \$/Deutschmark series for the years 1992, 1995, 1998, and 2001. None of these data are publicly available on the BIS website. Note further that the volume data mentioned above does not coincide with the temporal period in our sample. For details see <http://www.bis.org/statistics/index.htm> and BIS (2001). We encountered similar data constraint problems when we checked the IMF, World Bank and OECD websites.

<sup>19</sup> The number of potential coalitions in a bargaining situation is  $2^p - 1$ , where  $p$  is the number of legislative parties.

<sup>20</sup> Lanny Martin kindly provided us with data for the measures of uncertainty associated with government formation and future policy, as used here.

$i$ , coalition  $j$  will form is given by,

$$P_{ij} = \frac{\exp(\mathbf{x}'_{ij}\boldsymbol{\beta})}{\sum_{k=0}^{m_i} \exp(\mathbf{x}'_{ij}\boldsymbol{\beta})}, \quad (12)$$

where  $m_i$  is the number of potential coalitions in the bargaining situation,  $\mathbf{x}'_{ij}$  is a matrix of coalition attributes,<sup>21</sup> and  $\boldsymbol{\beta}$  is a vector of coefficient estimates associated with these attributes. All together, the data set we used, consists of 108 bargaining situations and 18,057 potential coalitions.<sup>22</sup> Once we obtained these probabilities, we constructed our index of bargaining uncertainty, which we refer to as *Effective Number of Coalitions* (ENC). The ENC index (for a given bargaining situation  $i$ ), is computed as follows:

$$ENC_i = \frac{1}{\sum P_{ij}^2}, \quad (13)$$

where  $P_{ij}$  represents the probability that coalition  $j$  in bargaining situation  $i$  will form. The ENC index is constructed along the same lines as the effective number of parties index developed by Laakso and Taagepera (1979) and possesses similar properties. In particular, ENC increases as the probability of formation becomes more dispersed among a greater number of coalitions, and it can range from a value of 1 to the full number of potential coalitions in a bargaining situation.<sup>23</sup> Thus, a larger value of ENC indicates greater uncertainty about which potential coalition will go on to form the government, therein allowing us to directly test Hypothesis 1. More substantively, our ENC variable captures *ex ante* uncertainty about (1) the ideological composition and ideological divisions within the future coalition government, (2) the relative bargaining strength of each party in the new coalition government, and (3) the number of parties that will comprise the coalition government. Our ENC measure is also designed to capture *ex ante* uncertainty that results from the formation of minority governments, which is a common phenomenon especially in Denmark.

Descriptive statistics of the ENC variable for the 10 European democracies in the 1970–1980s period—that are not reported to conserve space but are available on request—indicates that there is a fair degree of variation in bargaining uncertainty with respect to government formation in these countries. The mean value of ENC for Belgium (56.07) and Italy (12.91) indicates that uncertainty associated with government formation in these three countries is greater than uncertainty generated by coalition government formation in Austria (1.07), Ireland (2.52), Norway (0.51), and Sweden (2.77). This implies that uncertainty associated with coalition government formation in Italy is much higher compared with Norway and Sweden. The information from the mean ENC values is not surprising. For example, the bulk of the probability of coalition government formation in Austria, Germany, and

<sup>21</sup> The relevant attributes of a coalition include the following: whether it controls a lower-house majority, whether it is minimal winning, the number of parties that it contains, whether the largest legislative party is a member, whether the median legislative party is a member, the ideological divisions within the coalition, the ideological divisions within the opposition (relevant for minority coalitions only), whether the party of the previous prime minister is a member, whether the potential coalition is the incumbent government, whether an investiture rule is required (relevant for minority governments only), the degree of antisystem extremism within the coalition, whether the coalition is bound by an electoral pact, whether the coalition contains a party that is “very strong” or “merely strong” (à la Laver and Shepsle 1996), and whether a potential government containing such parties is a coalition or contains only a single party (Martin and Stevenson 2001).

<sup>22</sup> Martin and Stevenson (2001) restrict their analysis to situations in which no single party has won a majority of legislative seats.

<sup>23</sup> For example, given a three-party legislature, which would consist of seven potential coalitions, and probabilities for the potential coalitions equal to 0.01, 0.01, 0.94, 0.01, 0.01, 0.01, and 0.01, ENC would be equal to 1.13, reflecting the fact that almost all the probability is concentrated in one coalition. In contrast, in a three-party case where the probabilities for the seven coalitions are equal to 0.143, 0.143, 0.143, 0.143, 0.143, 0.143, and 0.143, ENC would be equal to 7 reflecting the fact that each of the seven potential coalitions is equally likely to form.

Sweden is normally concentrated on the left or in a particular coalition on the right. That is, bargaining uncertainty is usually low in these three countries thus explaining why the mean ENC value for these countries is low. Conversely, bargaining situations in Italy, for example, are known to be extremely complex and drag on for several months (Martin and Vanberg 2003). Hence, the mean ENC value for Italy is substantially higher. More interestingly, our ENC measure shows that uncertainty in Denmark and Norway—that have often had minority governments—is lower relative to Belgium and Italy.

The ENC measure also carefully captures temporal variation in uncertainty associated with government formation *within* countries. For instance, consider elections to the Belgian Parliament in December 1987. After the 1987 Belgian election, members from the Flemish and Wallon Socialist Parties and the Christian Social Party publicly indicated that they will form a coalition government and that the new government will adopt austerity measures to reduce inflation. Unfortunately, once bargaining over the composition of the future coalition government began, it became quite clear that the Flemish and Wallon Socialist parties were less inclined to become an integral part of the coalition government. As a result, agents in financial markets became increasingly concerned about the composition and policy objectives of the future coalition government (the *Financial Times*, January 2, 1988, Section 1:1). Given increased uncertainty about the composition of the future Belgian government, it is not surprising that the ENC value for Belgium December 1987 and January 1988—during which bargaining over the formation of the future government occurred—is almost 11.6% higher than the ENC value for February 1987 when uncertainty surrounding the composition of the future government considerably subsided.

Likewise, consider the 1989 Dutch election. The election results indicated that the standing Prime Minister, Christian Democrat Ruud Lubbers, would form a coalition government with the Labor Party (the *Guardian* (London), September 8, 1989). In a surprise, however, Lubbers was unable to make a deal with Labor and instead attempted to form a coalition government with the socialists who leaned toward higher government spending on welfare and higher agricultural subsidies (the *Economist*, November 4, 1989:62). The value of ENC for the Netherlands in September and October of 1989, during which hectic negotiations over the future government occurred, is thus almost 14% higher than the mean ENC value for this country. We introduce ENC in the conditional mean and variance equations of the GARCH model and expect that it will be negatively correlated with the mean and volatility of the exchange rate given the prediction in Hypothesis 1.

To test Hypothesis 2, we need a measure of uncertainty of trader expectations concerning government policy in situations of coalition bargaining. Following Martin (2004), we construct such a measure in several steps. First, for a given bargaining situation, we define the policy position of each potential coalition in that bargaining situation as the *weighted average* of the policy positions of the parties comprising the coalition.<sup>24</sup> To estimate party policy positions, we use the data set provided by Huber and Gabel (2000), which places parties on the left–right socioeconomic scale on the basis of the issue dimensions coded by the Manifestos Research Group (see Budge et al. 2001). Second, we compute the *expected* policy position of each potential coalition in the bargaining situation by multiplying its weighted-average position by the *ex ante* probability that the potential coalition will form (again using the predicted probabilities from the Martin and Stevenson 2001 analysis). Third, we sum the expected policy positions across all the potential coalitions in the bargaining situation. This constitutes our measure of *weighted-average*

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<sup>24</sup> The weights used are the legislative seat shares contributed to the government by each of the parties in the coalition. Thus, this measure takes into account the possibility that larger parties in the coalition, other things being equal, are probably better able to pull government policy toward their position than smaller parties.

*expected (government) policy* for the bargaining situation under consideration. Next, we define uncertainty of the weighted-average expected government policy using a measure of dispersion that is very similar to the standard descriptive statistic for population variance. Specifically, for any given bargaining situation, our measure of uncertainty of expected policy is

$$\sum_i p_i \left| pos_i - \sum_i p_i pos_i \right|, \quad (14)$$

where  $p_i$  is the probability that a potential coalition  $i$  will form and  $pos_i$  is the weighted-average policy position of coalition  $i$ . Thus, this measure takes the distance between the policy position of a potential coalition and expected government policy in a particular bargaining situation, discounts this distance by the probability that the potential coalition will form, and then sums the discounted deviations from expected policy across all the potential coalitions in the bargaining situation. We denote this measure as *Dispersion of Expected (government) Policy*, and we introduce it in the conditional mean and variance equations of the GARCH model. The measure has the desirable property that it becomes larger as the distribution of expected policies for the individual potential coalitions in the bargaining situation becomes more spread out around expected government policy. Following the prediction in Hypothesis 2, we expect that *Dispersion of Expected Policy* will be negatively correlated with the mean and volatility of the exchange rate.

The first main control variable that we introduce in the mean and variance equation of the GARCH model is *Formation Months*, which is a count of the number of months that a formation persists. It is assigned a value of zero during months when there is no formation, is coded 1 during the first month of a formation, 2 in the second month, and so on. We expect that *Formation Months* will be negatively correlated with the mean and volatility of the exchange rate because of the following reason: specifically, during months of government formation uncertainty about the potential coalition government that may form is likely to be high. As uncertainty about the coalition government formation that may form leads to lower volume and variability in the volume of trading (as predicted by our formal model), we would expect that both volume and variability of volume of trading will decrease during the months of government formation. This, in turn, will contribute to a lower mean and volatility of the exchange rate.

We add two other political controls to the conditional variance equation in the GARCH model. First, we introduce a dummy variable *European Monetary System (EMS)* to indicate the period after which a given country joined the EMS (this excludes Norway and Sweden).<sup>25</sup> For *EMS*, we would expect a decrease in currency volatility for countries that were subject to this mechanism (Leblang and Bernhard 2001:8). Second, we control for the expected partisan policy position of the governing coalition. As our measure of expected policy, we use the weighted-average expected partisan policy position for any given bargaining situation, which is defined as:  $\sum p_i pos_i$ . This measure, which we denote as *Expected Partisan Policy Position*, is one of the components used in the policy dispersion measure discussed above and defined formally in (14). Larger values of this variable indicate a greater likelihood that bargaining will result in the formation of a coalition supporting conservative economic policies, while smaller values indicate a greater likelihood that a leftist coalition will form.

Extant findings on the impact of partisanship on exchange rates are mixed. For instance, Clark and Hallerberg (2000) argue that there are no observable partisan

<sup>25</sup> Sweden joined the EMS on May 17, 1991 (and withdrew in 1992), which is after our temporal range ends for the DM/Swedish Krona series. Norway has not joined the EMS.

effects with regards to national monetary policy. Hence, according to Clark and Hallerberg (2000), the effect of partisanship on exchange rate volatility will be negligible. In contrast, Bearce (2003) and Freeman, Hays, and Stix (2000) find that partisanship does affect exchange rate volatility with left (right) parties associated with greater (lower) exchange rate variability, while Frieden (2002) finds partisan effects in the opposite direction. Given the mixed findings in the literature, we do not have a strong a priori expectation on how *Expected Partisan Policy Position* will affect exchange rate volatility in the GARCH model.<sup>26</sup>

With respect to economic control variables, we first include the variable *Output Volatility* in the conditional variance equation as Flood and Rose (1995:17–18)—who employ monthly exchange rate data for European countries (as we do)—find that volatility of output substantially and significantly affects exchange rate volatility. Given Flood and Rose’s (1995) finding, we use domestic industrial production indices from the countries that we examine as our measure of output.<sup>27</sup> We operationalized our measure of *Output Volatility* for each bilateral exchange rate series as follows: First, we constructed a measure of *Output Level*, which is defined as the ratio of (the natural logarithms of) German output to foreign output for each of the 10 exchange rate series that we examine. We then computed the sample standard deviation of the first difference of *Output Level* for each of the 9 exchange rate series—this measure constitutes the *Output Volatility* variable for each exchange rate sample. Following Flood and Rose (1995), we expect that *Output Volatility* will have a positive effect on exchange rate volatility.

The Flexible and Sticky–Price model predicts that in an exchange rate dyad between, say, country *i* and *j*’s currency, the difference between *i* and *j*’s (i) inflation rate, (ii) industrial output, and (iii) interest rate will affect the mean exchange rate level of *i* and *j*. Based on the predictions from the Flexible and Sticky–Price monetary models (see Taylor 1995:21–22), we introduce *Inflation* and *Interest Rate Differential* in the mean equation where it is expected to have a positive effect. Likewise, we introduce *Output Differential* in the mean equation, where it is expected to have a negative effect on the mean exchange rate level. As Rose (1996) finds that macroeconomic fundamentals have a weak but significant effect on exchange rate volatility, we also include *Inflation* and *Interest Rate Differential* in the conditional variance equation of the GARCH models. Data on macroeconomic fundamentals for the 9 European democracies are taken from the IMF’s *Government Financial Statistics* CD-Rom. We do not report summary statistics for our variables to save space, but these are available on request from the authors.

### Findings and Analysis

We first analyze the effects of the key variables of interest on exchange rate volatility in all nine countries as reported in Tables 2 and 3. We then discuss the results obtained for the mean exchange rate level.

The estimate of ENC in the conditional variance equations of the GARCH models—see Tables 2 and 3—is negative and statistically significant for all exchange rate series except for Ireland. This finding supports the prediction in Hypothesis 1 from our formal model. The estimated coefficient of ENC in the conditional variance equation is negative and statistically significant for (i) Austria ( $-0.0189$ ,  $t \approx 2.82$ ), (ii) Belgium ( $-0.0058$ ,  $t \approx 3.05$ ), (iii) Denmark ( $-0.0037$ ,  $t \approx 2.31$ ), (iv)

<sup>26</sup> The correlation between *Expected Partisan Policy Position* and *Dispersion of Expected Policy* is statistically insignificant and extremely low (0.007), thus mitigating concerns of collinearity between these two variables.

<sup>27</sup> Data for this variable are taken from *Federal Reserve Statistical Release* and the Bureau of Labor Statistics; see [http://www.federalreserve.gov/Releases/G17/Hist/table1\\_2.htm](http://www.federalreserve.gov/Releases/G17/Hist/table1_2.htm) and <http://www.federalreserve.gov/releases/h15/data/m/fp1m.txt>

TABLE 2. Results from GARCH Models (DM Denominated Exchange Rates)

	<i>Austria</i>	<i>Belgium</i>	<i>Denmark</i>	<i>France</i>	<i>Ireland</i>	<i>Italy</i>
Mean equation						
AR(1)	0.1702*** (0.0401)	0.2104*** (0.0618)	0.2526*** (0.0749)	0.2167*** (0.0479)	0.1948*** (0.0523)	0.0832* (0.0347)
ENC	-0.0101 (0.0096)	-0.0050 (0.0047)	-0.0031 (0.0039)	-0.3164 (0.3034)	-0.0892 (0.5155)	-0.1412 (0.8811)
Dispersion in expected policy	-0.0295 (0.0604)	-0.0516 (0.0773)	-0.0295 (0.0762)	0.0284 (0.0433)	-0.0526 (0.0497)	1.012 (3.185)
Formation months	-0.0472 (0.1892)	-0.0749 (0.1056)	-0.0752 (0.1219)	0.1842 (0.1478)	0.1184 (0.3176)	-2.893 (2.774)
Interest rate differential	-0.0179 (0.1642)	-0.2131 (0.1845)	-0.4061 (0.3772)	-0.1544 (0.1239)	-0.2643 (0.1825)	-20.125 (27.052)
Inflation differential	1.143 (10.95)	4.610 (12.08)	2.235 (11.14)	0.6829 (1.491)	1.037 (7.08)	401.11 (722.50)
Output differential	0.1252 (0.6177)	0.1820 (0.7544)	0.1876 (0.7659)	0.1476 (0.4715)	0.2122 (0.509)	-30.05 (100.06)
Variance equation						
ARCH	0.3073*** (0.0051)	0.2126*** (0.0049)	0.2543*** (0.0049)	0.2314** (0.0115)	0.2286** (0.1132)	0.1062 (0.0991)
GARCH	0.2980*** (0.0972)	0.3790*** (0.1161)	0.4131*** (0.1274)	0.4126*** (0.0773)	0.4932*** (0.0746)	0.4062 (0.3282)
ENC	-0.0189*** (0.0067)	-0.0058*** (0.0019)	-0.0037*** (0.0016)	-0.0108*** (0.0031)	0.0118 (0.2543)	-12.27** (6.782)
Formation months	-0.0147 (0.0254)	-0.0361 (0.0215)	-0.0551 (0.0384)	0.1652 (0.2154)	0.1165 (0.2120)	-33.25 (68.45)
Dispersion in expected policy	-0.012 (0.0319)	-0.0022 (0.0293)	-0.0019 (0.0387)	-0.0414 (0.0367)	-0.0321 (0.0421)	-40.031 (175.02)
Expected partisan policy position	-0.0039* (0.0019)	-0.0038** (0.0016)	-0.0036 (0.0029)	-0.0037* (0.0018)	-0.0035 (0.0024)	-5.143** (2.171)
Output volatility	0.1212 (0.1485)	-0.0173 (0.0193)	0.0131 (0.0089)	0.0065** (0.0032)	0.0126 (0.0109)	4.182 (7.351)
Inflation differential	0.0198 (0.0254)	0.0416 (0.0297)	0.0319 (0.0354)	0.7546 (1.81)	0.0289 (0.0256)	0.0284 (0.0433)
Interest rate differential	0.0031 (0.0040)	0.0065 (0.0055)	0.0078 (0.0082)	0.1255 (0.4350)	0.0059 (0.0060)	-87.250 (207.65)
EMS	-0.1129** (0.0514)	-0.6053*** (0.0437)	-0.5143*** (0.0552)	-0.2078*** (0.0544)	-0.2430*** (0.0571)	-82.73 (213.48)
Constant	0.5071*** (0.0121)	0.7138*** (0.0060)	0.4152*** (0.0069)	0.5134*** (0.0039)	0.4211*** (0.0168)	621.70 (191.22)
Diagnostics						
Ljung–box	15.689	12.237	12.232	11.348	10.998	15.434
Ljung–box squared	20.064	19.458	19.778	14.259	18.765	27.161
Kurtosis	1.105	1.083	1.146	1.092	0.945	1.128
<i>N</i>	126	228	164	211	217	201

*Notes:* Cell entries are maximum-likelihood estimates. Coefficient of constant in mean equation not reported to conserve space. Numbers in parentheses are Bollerslev–Wooldridge semi-robust standard errors.

\*\*\*denote significance at the 1%, 5%, and 10% levels respectively.

EMS, European Monetary System; GARCH, generalized autoregressive conditional heteroskedasticity; ARCH, autoregressive conditional heteroscedasticity; ENC, effective number of coalitions.

France (-0.0108,  $t \approx 3.48$ ), (v) Italy (-12.27,  $t \approx 1.86$ ), (vi) Luxembourg (-0.0182,  $t \approx 2.27$ ), (vii) the Netherlands (-0.0592,  $t \approx 8.45$ ), (viii) Norway (-0.009,  $t \approx 4.5$ ),<sup>28</sup> and (ix) Sweden (-0.0175,  $t \approx 2.21$ ).

<sup>28</sup> Residual diagnostics from the DM/Norwegian Krona series reveals that serial correlation in the mean and variances are corrected for only if we include an AR(1) and MA(1) term in the GARCH model. Hence, we included an MA(1) term for the DM/Krona series.

TABLE 3. Results from GARCH Models (Continued) (DM Denominated Exchange Rates)

	Luxembourg	The Netherlands	Norway	Sweden
Mean equation				
AR(1)	0.3385*** (0.0774)	0.2064*** (0.0492)	0.0479 (0.5231)	0.3122*** (0.0709)
MA(1)			0.0636 (0.7778)	
ENC	-0.1245 (0.1438)	-0.1120 (0.5641)	-0.0053 (0.0126)	-0.0998 (0.1541)
Dispersion in expected policy	0.0350 (0.0308)	-0.0108 (0.2477)	-0.0018 (0.0176)	0.0253 (0.0441)
Formation months	0.1361 (0.0904)	-0.1917 (0.2164)	-0.1655 (0.1023)	0.1289 (0.0872)
Interest rate differential	-0.0271 (0.0213)	-0.0519 (0.1253)	-0.0571 (0.0451)	-0.0153 (0.0240)
Inflation differential	-0.4508 (1.081)	0.7443 (3.614)	-1.068 (1.212)	-0.4032 (1.192)
Output differential	0.1792 (0.2140)	0.1825 (1.077)	0.0341 (0.2519)	0.1644 (0.2123)
Variance equation				
ARCH	0.0289*** (0.0945)	0.1233*** (0.0427)	0.0792** (0.0383)	0.1147*** (0.0823)
GARCH	0.5438*** (0.1182)	0.3744*** (0.0520)	0.3652*** (0.2120)	0.4320*** (0.1097)
ENC	-0.0182** (0.0080)	-0.0592*** (0.0070)	-0.009*** (0.002)	-0.0175** (0.0079)
Formation months	-0.0025 (0.0170)	-0.0053 (0.1288)	-0.0210 (0.0637)	-0.0036 (0.0151)
Dispersion in expected policy	-0.0051 (0.0044)	-0.4721 (0.6450)	-0.0019 (0.0056)	-0.0048 (0.0036)
Expected partisan policy position	-0.0060** (0.0028)	-0.0144 (0.0157)	-0.0001 (0.0015)	-0.0071** (0.0029)
Inflation differential	0.0016 (0.0402)	0.0307 (0.0252)	0.0126 (0.0265)	0.0114 (0.0330)
Interest rate differential	-0.0019 (0.0512)	0.0139 (0.0405)	0.0427 (0.0436)	0.0162 (0.7739)
Output volatility	7.07e - 05 (9.41e - 05)	0.0025 (0.0643)	-1.11e - 05 (1.05e - 05)	6.62e - 05 (8.17e - 05)
EMS	-0.0139 (0.0094)	-0.1218*** (0.0479)	—	—
Constant	-0.0033 (0.0064)	0.3011*** (0.0527)	0.0741*** (0.0050)	-0.0032 (0.0065)
Ljung-Box	11.029	12.432	11.503	10.962
Ljung-Box Squared	18.954	19.248	20.517	17.655
Kurtosis	1.017	1.216	1.022	0.916
N	167	148	228	181

Notes: Cell entries are maximum-likelihood estimates. Coefficient of constant in mean equation not reported to conserve space. Numbers in parentheses are Bollerslev-Wooldridge semi-robust standard errors.

\*\*\*\*\*denotes significance at the 1%, 5%, and 10%, respectively.

EMS, European Monetary System; GARCH, generalized autoregressive conditional heteroskedasticity; ARCH, autoregressive conditional heteroscedasticity; ENC, effective number of coalitions.

We also derived the *substantive* effect of ENC by examining the extent to which a one standard deviation increase in ENC affects exchange rate volatility when the remaining variables in the conditional variance equation are set at their mean. We find that a one standard deviation increase in ENC decreases volatility of the (i) DM/Austrian Schilling by 3.3%, (ii) DM/Belgian Franc by a modest 0.6%, (iii) DM/Danish Krone by 4%, (iv) DM/French Franc by 4%, (v) DM/Lira by 3.7%, (vi) DM/Lux-

embourg Franc by 5%, (vii) DM/Guilder by 4.1%, (viii) DM/ Norwegian Krone by 1%, and (ix) DM/ Swedish Krona by 2.6%. Our finding that uncertainty about the future potential coalition government has a significant negative effect on exchange rate volatility is important for two reasons.

First, it indicates that a more refined and accurate measure of uncertainty associated with government formation in parliamentary democracies provides a strikingly different result compared with those that are claimed and presented in the extant literature. Second, the findings on ENC arguably confirm our theoretical intuition that political uncertainty about the potential coalition government that may form gives incentives to traders to reduce the volume of speculative trading, therein helping to reduce volatility.

Our findings on ENC discussed above raise two questions. First, why does the substantive effect of ENC differ significantly across the nine exchange rate series being analyzed here? It is plausible that variation in substantive effects across different exchange rates could have resulted from variability in traders' response to uncertainty stemming from the politics of government formation. For example, traders who trade the French Franc could be reacting more aggressively and with more alacrity to uncertainty about government formation relative to those who trade the Norwegian Kroner. Furthermore, it is possible that uncertainty about the future government could be having stronger effects on the volume of the French Franc traded relative to the volume of the traded Kroner. Differences in trading volume could also engender variation in substantive effects. Second, why do our results on the correlation between uncertainty associated with government formation and exchange rate volatility differ from those in the existing literature? We address this question in some detail at the end of this section.

For now, note that unlike ENC, the estimate of *Dispersion of Expected Policy* in the conditional variance equations provides weak support for the prediction in Hypothesis 2. The estimate of this variable in the conditional variance equation of all the GARCH models has the predicted negative sign but is not statistically significant in any model. While the lack of statistical significance is disappointing, it does partially support our theoretical claim that the rational response of currency traders to uncertainty over future policies leads to a decline—and not increase—in exchange rate volatility.

We obtain mixed results for the political controls in the conditional variance equation of the estimated GARCH (1, 1) models. The estimate of EMS in the variance equation has the predicted negative sign and is highly statistically significant in most of the exchange rate series examined for our empirical analysis. This result confirms existing studies, which find that EMS has a substantive negative effect on exchange rate volatility (see Bernhard and Leblang 2002a). The estimate of *Formation Months* has the predicted negative sign, but is not statistically significant in any of the exchange rate series that we analyze. While the negative estimate of *Formation Months* confirms our intuition that trading volume drops during periods of government formation, which contributes to a decline in currency volatility, the insignificance of this variable indicates that the substantive effect of *Formation Months* on currency volatility is weak.

The estimate of *Expected Partisan Policy Position* is negative in the variance equation of all the GARCH models and statistically significant in six exchange rate series. This result supports the Bearce (2003) and Freeman, Hays, and Stix (2000) findings that right-wing (left-wing) governments are associated with lower (higher) exchange rate volatility, but contradicts Frieden's (2002) finding that partisanship has the opposite effect. Moreover, our finding that partisanship has a negative effect on currency volatility suggests, in contrast to Clark and Hallerberg's (2000) claim, that currency markets expect that government partisanship will influence monetary policy. We do not discuss why our results on the impact of partisanship on exchange rate variability are different from those obtained by Frieden (2002) and Clark and Hallerberg (2000), but this is an issue-area which requires further research.

Turning to the economic controls, we find that the estimate of *Output Volatility* in the variance equation is for the most part insignificant. *Interest Rate* and *Inflation Differential* is statistically insignificant in the conditional variance equation in all the empirical models in Tables 2 and 3. The aforementioned results confirm Meese and Rogoff's (1983b) claim that macroeconomic fundamentals do not have a significant effect on exchange rate volatility.

We now briefly describe the parameter estimates obtained in the conditional mean equation for each exchange rate series. Observe that the estimate of ENC in the conditional mean equation is negative, but statistically insignificant for all the exchange rate series. The estimate of *Dispersion of Expected Policy* and *Formation Months* is insignificant in all the exchange rate series examined here. The estimates of the economic variables in the mean equation of all the GARCH models are also largely insignificant and thus inconsistent with the predictions of the Flexible and Sticky-Price monetary models.

Why are the estimates of the political and economic variables largely insignificant in the mean equation? First, given the efficient markets hypothesis—that the mean level of exchange rates at any given month reflects all available information and is merely a function of past prices—the insignificant results that we obtain in the mean equation are not striking. Second, the coefficient of the AR(1) parameter in the mean equation is positive and significant in almost all the GARCH models in Tables 2 and 3. This finding further confirms that the mean exchange rate level in all our samples is largely influenced by past prices. Third, our finding that macroeconomic variables have insignificant effects in the mean equation confirms existing empirical studies by economists who find that macroeconomic fundamentals do not have a statistically significant effect on mean returns of currency markets (Taylor 1995:33–34).

We provide three reasons to explain why our empirical results on the correlation between uncertainty about the potential coalition government that may form and volatility differ from those in the existing literature. First, in contrast to the literature that uses dummy variables to capture uncertainty over government formation and dissolution (Leblang and Bernhard 2001; Bernhard and Leblang 2002a), we use a *continuous* measure of uncertainty associated with coalition government formation that captures the multidimensional nature of *ex ante* uncertainty that economic agents face in situations of coalition bargaining. Operationalizing uncertainty as a continuous variable provides us with a richer measure of the true degree of temporal variation in uncertainty about the potential coalition government formation that may form within and across advanced parliamentary democracies with a history of coalition governments. This is critical because we believe that a richer operationalization of the temporal dynamics of uncertainty allows us to capture more accurately the correlation between uncertainty associated with government formation and exchange rate volatility, which is negative and not positive.

Second, in contrast to existing empirical models of the effects of political uncertainty on FOREX markets, our GARCH models are well specified and theoretically driven. For instance, unlike current empirical models that do not explicitly theorize or capture how political *and* economic variables affect the mean exchange rate level,<sup>29</sup> we not only introduce key economic controls in the mean equation of our GARCH model, but also political variables that are informed by a well-grounded theory. It is possible that the inclusion of political variables and economic controls in the mean equation of the GARCH models led to results that are different from those in the current literature. Third, we control for important macroeconomic variables in the mean and variance equation of the GARCH model, which is typically ignored in the existing empirical literature on democratic politics and finan-

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<sup>29</sup> See, for example, Freeman, Hays, and Stix (2000) and Leblang and Bernhard (2001).

cial markets. It is plausible that explicitly controlling for various macroeconomic variables may have led to results that differ from those posited in the literature.

In results not reported here in order to conserve space, we checked the robustness of the findings in Tables 2 and 3 by including in the variance equations of each GARCH model the following economic variables: the level of bilateral trade as a ratio of Germany's GDP and the degree of Central Bank independence in each country. Additional variables such as the lag of ENC, the lag of *Dispersion of Expected Policy*, the lag of *Formation Period* and a dummy for election months in the nine European democracies were included in the variance equation in each GARCH model. In no case did inclusion of any of these variables significantly—substantively or statistically—alter the results reported in Tables 2 and 3.<sup>30</sup> We also tested our two hypotheses for each European exchange rate series with respect to the U.S. dollar. The results reported in Tables 2 and 3 remain robust when the exchange rate series are denominated in U.S. dollars.

Turning to diagnostics, we find that in all the GARCH models in Tables 2 and 3, the Ljung-Box statistics reject the null of remaining serial correlation in levels (Ljung-Box statistic) as well as in variances (Ljung-Box squared). This indicates that the empirical models in Tables 2 and 3 are well specified. In almost all the empirical models, the estimate of the ARCH term ( $\varepsilon_{t-1}^2$ ) and the GARCH term ( $\sigma_{t-1}^2$ ) is statistically significant. This means that the shock  $\varepsilon_t$  from the prior period ( $\varepsilon_{t-1}^2$ ) and the conditional variance from time  $t-1$  significantly affects the conditional variance at time  $t$ , which implies that equation (2) in our formal model holds true in the data. We implemented the Geweke and Porter-Hudak (1983) log periodogram estimator, also known as the GPH test, to check whether the exchange rate series that we examine are fractionally integrated.<sup>31</sup>  $p$ -values from the GPH test for each of the 10 differenced exchange rate series indicate that they are not fractionally integrated. This indicates that the findings reported in Table 2 are not spurious. Finally, we tested for multicollinearity problems in all the GARCH models by using variance inflation factor (VIF) tests. The tests reveal that multicollinearity is not a problem in any of the empirical models in Tables 2 and 3.<sup>32</sup>

### Implications and Conclusions

This study contributes to the study of democratic politics and FOREX markets in two main ways. First, unlike the extant literature, we construct a formal model that provides a well specified causal theory of how currency traders rationally respond to political uncertainty associated with the composition of the future coalition government formation and how this, in turn, affects the exchange rate process. Second, in contrast to existing studies, our empirical findings suggest that uncertainty about the potential coalition government that may form has a statistically significant *negative* impact on the volatility of exchange rates.

This article's central findings have numerous important implications. First, scholars argue that bargaining associated with coalition government formation hinders efficient policy making and may have negative welfare consequences on financial markets and the economy (Laver and Schofield 1990; Freeman, Hays, and Stix 2000;

<sup>30</sup> Results from specifications with these additional variables are available upon request.

<sup>31</sup> The GPH test provides an estimate of the fractional differencing parameter  $\hat{d}_{GPH} = \sum_{s=1}^m y_s \log(I_y(\xi_s)) / 2 \sum_{s=1}^m y_s^2$  where  $\omega_y(\xi_s) = 1 / \sqrt{2\pi m} \sum_{t=1}^m y_t e^{it\xi_s}$  is the discrete Fourier transform of the exchange rate series.  $p$ -values from the GPH test for each differenced exchange rate series for  $n = T^{0.45}$ ,  $n = T^{0.60}$ , and  $n = T^{0.55}$  indicate that they are not fractionally integrated.

<sup>32</sup> Chatterjee, Price, and Haidi (1999) suggests that multicollinearity exists if (i) the largest VIF is greater than 10; and (ii) the mean VIF is larger than 1. The mean and largest VIF values for all the GARCH models we estimated were substantially lesser than the threshold values mentioned above, thus indicating that multicollinearity is not a problem.

Bernhard and Leblang 2002a). This conclusion is premature. Based on our formal model, we argue that it is difficult to clearly ascertain the welfare consequences of the impact of uncertainty about government formation on currency markets. On the one hand, it appears that political uncertainty, at least with respect to the coalition governments that may form, surprisingly has positive welfare consequences in financial markets because it leads to lower exchange rate volatility. On the other hand, however, political uncertainty about the coalition government that might form may reduce the mean exchange rate level that can foster capital outflows from a particular country and thus have negative welfare consequences. This ambiguity indicates that we cannot draw simple inferences about the welfare consequences of uncertainty associated with government formation on currency markets.

Second, our study reveals that as traders automatically adjust and hedge against political uncertainty—which helps to minimize volatility—politicians ironically have low incentives to “calm” markets by reducing uncertainty associated with the formation of coalition governments. Indeed, it will be irrational for politicians that are involved in coalition bargaining to expend extra effort or undertake costs to stabilize financial markets given that traders adjust to political uncertainty engendered by government formation. While the aforementioned observation may sound trivial, it has far-reaching and perverse political implications.

First, it suggests that politicians involved in the coalition bargaining process are likely to bargain longer for ministerial portfolios in a future coalition government especially if they recognize that uncertainty associated with government formation does not have detrimental effects on exchange rates. Second, it suggests that politicians in countries with coalition governments may have low *economic* incentives to design institutions to minimize the frequency of coalition government formation given that the politics of coalition formation may not have costly consequences on FOREX markets and economic outcomes in general. Third, the economics literature on exchange rate dynamics ignores the impact of domestic political uncertainty on FOREX markets (see Taylor 1995). Our study joins a growing literature on the political economy of exchange rates showing that political variables including uncertainty about government formation has substantial effects on the volatility of exchange rates.

The research presented here can be extended in two main directions. First, the empirical domain might be broadened. It might be worthwhile to test the hypotheses from our model on exchange rates of developing countries with a parliamentary democracy and coalition governments. Extension to other FOREX markets such as the forward exchange swap market is a potentially fruitful direction. Whatever future directions prove fruitful, we hope that this study provides a step toward a more careful understanding of how political uncertainty associated with government formation and policy expectations affect exchange rate dynamics in advanced industrial parliamentary democracies.

## Appendix A

**Lemma 1:** *The optimal amount (volume) of the foreign currency that is bought by the buyer is*

$$k_{f,b}^* = \frac{a - (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t)}{(1 + \gamma_{k_{f,b}} \sigma_\psi^2 + \sigma_m^2)}. \quad (\text{A.1})$$

*Proof:* Substituting  $E(\pi_B|S_t)$  and  $Var(\pi_B|S_t)$  into the buyer’s utility function leads to  $\left(a - (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t)k_{f,b} - \frac{(k_{f,b})^2}{2}\right) - \left((\alpha\sigma_\psi^2 + \sigma_m^2)(k_{f,b})^2\right)$ . Simplifying the aforementioned expression and maximizing the buyer’s utility function with respect to

$k_{f,b}$  yields  $k_{f,b}^*$ , as stated in (A.1). To obtain an economically meaningful (positive) solution for  $k_{f,b}^*$ , we let  $a > (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t)$ . ■

**Lemma 2:** *The optimal amount (volume) of the foreign currency that is sold by the seller is*

$$k_{f,s}^* = \frac{\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t - c}{(1 + \gamma_{k_{f,b}} \sigma_\psi^2 + \sigma_m^2)} \text{ where } \bar{e} + \phi\varepsilon_{t-1} + \alpha S_t > c. \quad (\text{A.2})$$

Substituting equations  $E(\pi_s|S_t)$  and  $Var(\pi_s|S_t)$  into the seller's utility function leads to  $\left((\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t - c)k_{f,s} - \frac{(k_{f,s})^2}{2}\right) - \left((\alpha\sigma_\psi^2 + \sigma_m^2)(k_{f,s})^2\right)$ ; maximizing this expression with respect to  $k_{f,s}$  yields  $k_{f,s}^*$ , as stated in (A.2). ■

*Proof of Proposition 1:* Since the signal is a function of  $g_t$  and  $\psi_t$ , the stochastic component of the signal is  $\alpha S_t = \frac{Cov(g_t, S_t)}{Var(S_t)} = \frac{\sigma_g^2}{\sigma_g^2 + \sigma_\psi^2} S_t$ , that is, the variance decomposition of  $\alpha S_t$  in the denominator is  $Var(S_t) = \sigma_g^2 + \sigma_\psi^2$ , while in the numerator, it is  $Cov(g_t, S_t) = \sigma_g^2$ . Given  $\alpha S_t$ , we obtain

$$\frac{\partial k_{f,b}^*}{\partial \sigma_g^2} = -\frac{\sigma_\psi^2}{(\sigma_g^2 + \sigma_\psi^2)} \left\{ \frac{(1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2))S_t + \gamma\sigma_\psi^2(a - (\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t))}{(1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2))^2} \right\} < 0$$

and

$$\frac{\partial k_{f,s}^*}{\partial \sigma_g^2} = -\frac{\sigma_\psi^2}{(\sigma_g^2 + \sigma_\psi^2)} \left\{ \frac{(1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2))S_t + \gamma\sigma_\psi^2((\bar{e} + \phi\varepsilon_{t-1} + \alpha S_t) - c)}{(1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2))^2} \right\} < 0 \quad \forall S_t$$

$$> 0.$$

As  $\hat{e}_t$  is linear in  $k_{f,b}^*$  and  $k_{f,s}^*$  and because  $\frac{\partial k_{f,b}^*}{\partial \sigma_g^2} < 0$ ,  $\frac{\partial k_{f,s}^*}{\partial \sigma_g^2} < 0$ ,  $\hat{e}_t$  decreases when  $\sigma_g^2$  increases.

$$\frac{\partial Var(k_{f,b}^*)}{\partial \sigma_g^2} = \frac{\partial Var(k_{f,s}^*)}{\partial \sigma_g^2} \equiv \left\{ \frac{-\left(\alpha(2 + \alpha)(1 + \alpha\sigma_m^2) + \gamma\alpha\sigma_\psi^2\right)}{(1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2))^3} \right\} < 0.$$

As  $Var(\hat{e}_t)$  is linear in  $Var(k_{f,b}^*)$  and  $Var(k_{f,s}^*)$  and  $\frac{\partial Var(k_{f,b}^*)}{\partial \sigma_g^2} < 0$ ,  $\frac{\partial Var(k_{f,s}^*)}{\partial \sigma_g^2} < 0$ ,  $Var(\hat{e}_t)$  decreases as  $\sigma_g^2$  increases. ■

*Proof of Proposition 2:*

$$\frac{\partial Var(k_{f,b})}{\partial \sigma_\psi^2} = \frac{\partial Var(k_{f,s})}{\partial \sigma_\psi^2} \equiv \left\{ \frac{\sigma_g^2 \left( (1 + \gamma(\sigma_\psi^2 \alpha + \sigma_m^2)) \frac{\alpha}{\sigma_g^2 + \sigma_\psi^2} + 2\gamma\alpha^3 \right)}{-\gamma(\sigma_\psi^2 \alpha + \sigma_m^2)^3} \right\} < 0.$$

As  $Var(\hat{e}_t)$  is linear in  $Var(k_{f,b})$  and  $Var(k_{f,s})$  and because  $\frac{\partial Var(k_{f,b})}{\partial \sigma_\psi^2} = \frac{\partial Var(k_{f,s})}{\partial \sigma_\psi^2} < 0$ , it follows that  $Var(\hat{e}_t)$  decreases when  $\sigma_\psi^2$  increases. Differentiating  $k_{f,b}^*$  and  $k_{f,s}^*$  with respect to  $\sigma_\psi^2$  yields

$$\frac{\partial k_{f,b}^*}{\partial \sigma_{\psi}^2} \equiv - \left\{ \frac{\sigma_g^2 \left( (1 + \gamma(\sigma_{\psi}^2 \alpha + \sigma_m^2)) \frac{\alpha}{\sigma_g^2 + \sigma_{\psi}^2} S_t - \gamma \alpha^2 (a - (\bar{e} + \phi \varepsilon_{t-1} + \alpha S_t)) \right)}{(1 + \gamma(\sigma_{\psi}^2 \alpha + \sigma_m^2))^2} \right\} < 0,$$

$$\frac{\partial k_{f,s}^*}{\partial \sigma_{\psi}^2} \equiv - \left\{ \frac{\sigma_g^2 \left( (1 + \gamma(\sigma_{\psi}^2 \alpha + \sigma_m^2)) \frac{\alpha}{\sigma_g^2 + \sigma_{\psi}^2} S_t - \alpha^2 ((\bar{e} + \phi \varepsilon_{t-1} + \alpha S_t) - c) \right)}{(1 + \gamma(\sigma_{\psi}^2 \alpha + \sigma_m^2))^2} \right\} < 0.$$

As  $\hat{e}_t$  is linear in  $k_{f,b}^*$ , these derivatives imply that  $\hat{e}_t$  decreases when  $\sigma_{\psi}^2$  increases. ■

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