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Exchange Rate Models Are Not as Bad as You Think

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There appears to be a consensus among researchers in exchange-rate economics that the standard models that relate exchange rates to monetary variables—prices, interest rates, and so forth—are off the mark. For example, Sarno and Taylor (2002, pp. 136–37) state, "Overall, the conclusion emerges that, although the theory of exchange rate determination has produced a number of plausible models, empirical work on exchange rates still has not produced models that are sufficiently satisfactory to be considered reliable and robust. In particular, although empirical exchange rate models occasionally generate apparently satisfactory explanatory power in-sample, they generally fail badly in out-of-sample forecasting tests in the sense that they fail to outperform a random walk." Bacchetta and van Wincoop (2006, p. 552) observe, "The poor explanatory power of existing theories of the nominal exchange rate is most likely the major weaknesses of international macroeconomics. Richard A. Meese and Kenneth Rogoff (1983a) and the subsequent literature have found that a random walk predicts exchange rates better than macroeconomic models in the short run." Evans and Lyons (2002, pp. 170–71) assert, "Macroeconomic models of exchange rates perform poorly at frequencies higher than one year. Indeed, the explanatory power of these models is essentially zero (Meese and Rogoff 1983a, Meese 1990). In the words of Frankel and Rose (1995, p. 1704), this negative result has had a 'pessimistic effect on the field of empirical exchange rate modeling in particular and international finance in general.' The pessimistic effect has been with us 20 years."

We present evidence that exchange rate models are not so bad after all. We approach the problem from several angles, but all of the approaches are linked by the observation that short-run movements in exchange rates are primarily determined by changes in expectations—exactly as

the standard models say. We begin in section 1 by demonstrating that standard models imply near-random walk behavior in exchange rates, so that their power to "beat the random walk" in out-of-sample forecasts is low. We then offer various alternative means for evaluating exchange rate models: (1) using in-sample fit of the models, but highlighting the role of endogeneity of monetary policy in explaining nominal and real exchange rate behavior (section 2); (2) examining whether exchange rates incorporate news that helps to predict the future macroeconomic fundamentals, as implied by the models (section 3); (3) reexamining the question of whether the models can account for the volatility of exchange rates (also section 3); (4) reviewing the recent literature that has examined the response of exchange rates to announcements of macroeconomic news (section 4); (5) presenting estimates of the model in which expectations of fundamentals are drawn from survey data (also section 4); and (6) demonstrating that the predictive power of the models can be greatly increased by using panel techniques and forecasting exchange rates at longer horizons (section 5). Conclusions are in section 6.

We begin by examining the theorem in Engel and West (2005) that demonstrates that under plausible assumptions, the models actually imply that the exchange rate should nearly follow a random walk. Therefore, it should not be surprising that the exchange rate models cannot provide better forecasts than the random walk model. As we will elucidate, the key insight behind the theorem is that current economic fundamentals have relatively little weight in determining the exchange rate in standard models. Much greater weight is put on expectations of future fundamentals, even fundamentals several years into the future.

We elaborate on the implications of the theorem. We show that in a standard parameterization of the famous Dornbusch "overshooting" model, the exchange rate nearly follows a random walk. We make this characterization in spite of the fact that the best-known feature of this model is that exchange rate changes should be predictable—that the exchange rate overshoots its long-run value in response to monetary shocks. We also argue that Meese and Rogoff's (1983a) exercise, in which the model forecasts use actual ex post (rather than forecasted) values of the fundamentals, is potentially flawed, because the out-of-sample fit of the models can be made arbitrarily worse or better by algebraic transformations of the model. Indeed, the out-of-sample fit of the standard models can be made much better (under the Meese-Rogoff methodology) if the models are written in a way that emphasizes the importance of expectations in determining exchange rates.

If we do not use the criterion of outperforming the random walk model in out-of-sample forecasting power, how then should we evaluate exchange rate models? We offer a number of alternatives.

First, we can look at the in-sample fit of the models. Most empirical monetary models of the 1970s and 1980s paid little attention to the endogeneity of monetary policy. But if exchange rates are primarily driven by expectations, then correctly modeling monetary policy is critical. Changes in current economic fundamentals, for example, may have a greater impact on exchange rates indirectly through the induced changes in expectations of monetary policy than through any direct channel. Engel and West (2006), Mark (2007), Clarida and Waldman (2006), and Molodtsova and Papell (2007) have explored the empirical performance of models based on Taylor rules for monetary policy. In a traditional flexible price model, an increase in current inflation depreciates the currency. But it is important to understand the policy reaction to higher inflation. In the new Taylor-rule models, higher inflation leads to an appreciation in inflation-targeting countries, because higher inflation induces expectations of tighter future monetary policy.

We review these models and also make note of the important result of Benigno (2004): the Taylor-rule models offer a potential solution to the "purchasing power parity" puzzle. In particular, the models offer the possibility that persistent real exchange rates do not require unrealistic assumptions about the stickiness of nominal price setting, and potentially de-link the two altogether.

Engel and West (2004, 2005) propose testing two implications of the present-value models. They emphasize that because we acknowledge that there are unobserved fundamentals (e.g., money demand shocks, risk premiums), the exchange rate may not be exactly the expected present value of observed fundamentals. But if exchange rates react to news about future economic fundamentals, then perhaps exchange rates can help forecast the (observed) fundamentals. If the observed fundamentals are the primary drivers of exchange rates, then the exchange rates should incorporate some useful information about future fundamentals. We verify this proposition using Granger causality tests. Engel and West (2004) also develop a technique for measuring the contribution of the present discounted sum of current and expected future observed fundamentals to the variance of changes in the exchange rate, which is valid even when the econometrician does not have the full information set that agents use in making forecasts. Here we find that the observed fundamentals can account for a relatively large fraction of ac-

tual exchange rate volatility, at least under some specifications of the models.

Standard tests of forward-looking models under rational expectations make the assumption that the sample distribution of ex post realizations of economic variables provides a good approximation of the distribution used by agents in making forecasts. But (as Rossi [2005] has emphasized) when agents are trying to forecast levels of variables that are driven by persistent or permanent shocks, the econometrician might get a very poor measure of the agents' probability distribution by using realized ex post values. The problem is enhanced when the data-generating process is subject to long lasting regime shifts (caused, for example, by changes in the monetary policy regime). For an economic variable such as the exchange rate—which is primarily driven by expectations—it might be useful to find alternative ways of measuring the effect of expectation changes.

Several recent papers (Andersen, Bollerslev, Diebold, and Vega [2002], Faust, Rogers, Wang, and Wright [2007], Clarida and Waldman [2006]) have looked at the effects of news announcements on exchange rates, using high-frequency data. We review these studies and argue that the response of exchange rates to news is precisely in line with the predictions of the Taylor-rule models.

We also provide new evidence of another sort. We directly measure expectations of inflation and output using surveys of professional forecasters. The particular survey we employ asks forecasters twice a year to provide predictions for inflation and output growth in a dozen advanced countries for the current year, each of the next five years, and an average for years 6–10. From these surveys, we are able to construct a present value of current and expected future fundamentals implied by the Taylor-rule model. We find strong confirmation of the model—higher output growth and higher inflation in the United States relative to the other countries leads to an appreciation of the dollar relative to the other currencies.

While we argue that theoretically the models may have low power to produce forecasts of changes in the exchange rate that have a lower mean-squared-error than the random walk model, we also explore ways of increasing the forecasting power. Mark and Sul (2001), Rapach and Wohar (2002), and Groen (2005) have used panel error-correction models to forecast exchange rates at long horizons (16 quarters, for example). We find that with the increased efficiency from panel estimation, and with the focus on longer horizons, the macroeconomic models consis-

tently provide forecasts of exchange rates that are superior to the "no change" forecast from the random walk model.

We find ourselves in the uncomfortable position of both pointing out that common formulations of monetary models imply that the models should have little power to produce better forecasts than a random walk, while at the same time finding more forecasting power for the models than many previous studies (though in line with the findings of the studies that employ panel techniques, cited earlier). There are two possible resolutions to this conflict. First, in section 5, we demonstrate in an example that if the foreign exchange risk premium—whose behavior is not well understood, and which is not observable to the econometrician—is stationary with an innovation variance that is relatively small (compared to variances in innovations of standard "observed" fundamentals), then the models might have predictive power relative to the random walk at long horizons in the error-correction framework. So it may be that indeed, using panel techniques, we have confirmed the usefulness of the models in forecasting at horizons of 16 quarters. The other resolution to the conflict is that our prediction results might prove fleeting: history has shown that models that seem to fit well over some time periods end up not holding up as the sample extends. That might hold true for forecasting power as well. While it is encouraging that our forecasting results confirm the findings of Mark and Sul (2001) on an extended sample, these results may ultimately prove not to be robust. But the Engel and West (2005) theorem tells us that out-of-sample prediction power relative to a random walk is not a reliable gauge to judge exchange rate models.

1 Present-Value Models and Random Walks

Let s_t denote the log of the exchange rate, measured as the log of the domestic currency price of foreign currency. Thus, depreciation of the currency implies an increase in s_t. Consider models of the exchange rate that relate the value of the currency to economic fundamentals, and to the expected future exchange rate:

$$s_t = (1 - b)\mathbf{a}_1'\mathbf{x}_t + b\mathbf{a}_2'\mathbf{x}_t + bE_t s_{t+1}, \quad 0 < b < 1. \tag{1}$$

Exchange-rate behavior is ultimately driven by $\mathbf{x}_t$, a vector of economic "fundamentals." Many familiar exchange rate models based on macroeconomic fundamentals take this general form, as subsequent examples will demonstrate.

Equation (1) is an expectational difference equation with a no-bubbles forward solution given by:

$$s_t = (1 - b)E_t\left(\sum_{j=0}^{\infty} b^j \mathbf{a}_1' \mathbf{x}_{t+j}\right) + bE_t\left(\sum_{j=0}^{\infty} b^j \mathbf{a}_2' \mathbf{x}_{t+j}\right). \tag{2}$$

The log of the exchange rate is determined as the expected present discounted value of current and future fundamentals. As in many models of other asset prices, if the discount factor is large (close to 1), expected future fundamentals matter a lot more than the current value of the fundamental. For example, if the fundamentals were expected to change between period t and $t + 1$ to new permanent values, $\tilde{\mathbf{x}}_{t+1}$, the exchange rate would be a weighted average of the current and the future fundamental—but with much more weight placed on the future fundamental:

$$s_t = (1 - b)\left(\mathbf{a}_1' \mathbf{x}_t + \frac{b}{1-b}\mathbf{a}_2' \mathbf{x}_t\right) + b\left(\mathbf{a}_1' \tilde{\mathbf{x}}_{t+1} + \frac{b}{1-b}\mathbf{a}_2' \tilde{\mathbf{x}}_{t+1}\right).$$

That is, in a model such as this, when fundamentals are very persistent we can say that the exchange rate is primarily determined by the expected future path of the fundamentals, with little weight given to the current fundamental.

The monetary exchange rate models of the 1970s and 1980s take the form given by equation (1). They are based on a Cagan-style money demand model. For the home country we can write:

$$m_t - p_t = \alpha + \gamma y_t - \lambda i_t + v_t, \tag{3}$$

where m_t is the log of the money supply, p_t is the log of the home consumer price level, y_t is the log of output, i_t is the home interest rate (in levels), and v_t is a stochastic shift term. Defining the real exchange rate as $q_t = s_t + p_t^* - p_t$, and assuming the foreign money demand equation has the same parameters as the home, we can write:

$$m_t - m_t^* - s_t + q_t = \gamma(y_t - y_t^*) - \lambda(i_t - i_t^*) + v_t - v_t^*. \tag{4}$$

Foreign variables are denoted with a*.

Now we will introduce the relationship:

$$i_t - i_t^* = E_t s_{t+1} - s_t + \rho_t. \tag{5}$$

This relationship defines ρ_t, the deviation from uncovered interest parity. As is well known, a vast empirical literature has rejected the hypothesis that $\rho_t = 0$. But so far there is no consensus on a model for ρ_t. Perhaps it is a risk premium, a short-run deviation from rational expec-

tations, or some other market imperfection. It is possible that movements in ρ_t are important in explaining exchange rate movements (as Obstfeld and Rogoff [2003] have suggested), but we do not explore that avenue in this paper. We treat ρ_t as an unobserved fundamental—an economic variable that might drive the exchange rate, but a variable for which we do not have direct observations. The money-demand shifts, v_t and v_t^*, are also treated as unobserved fundamentals.

When we combine equations (4) and (5), we get an equation that takes the form of equation (1):

$$s_t = \frac{1}{1+\lambda}[m_t - m_t^* + q_t - \gamma(y_t - y_t^*) - (v_t - v_t^*)] + \frac{\lambda}{1+\lambda}\rho_t + \frac{\lambda}{1+\lambda}E_t s_{t+1}. \tag{6}$$

In this case, the discount factor, b, from equation (1) corresponds to $\lambda/(1 + \lambda)$ in equation (6). The linear combination of fundamentals $\mathbf{a}_1'\mathbf{x}_t$ is given by $m_t - m_t^* + q_t - \gamma(y_t - y_t^*) - (v_t - v_t^*)$, while ρ_t corresponds to $\mathbf{a}_2'\mathbf{x}_t$. The no-bubbles forward solution to equation (6), then, is:

$$s_t = \frac{1}{1+\lambda}E_t\left\{\sum_{j=0}^{\infty}\left(\frac{\lambda}{1+\lambda}\right)^j[m_{t+j} - m_{t+j}^* + q_{t+j} - \gamma(y_{t+j} - y_{t+j}^*) - (v_{t+j} - v_{t+j}^*)]\right\} + \frac{\lambda}{1+\lambda}E_t\left[\sum_{j=0}^{\infty}\left(\frac{\lambda}{1+\lambda}\right)^j\rho_{t+j}\right]. \tag{7}$$

Equation (7) is representative of the type of model that we contend is "better than you think." That is, it is a rational-expectations model based on macroeconomic fundamentals. It is the traditional models that were explored in depth in the 1970s and 1980s. Moreover, a model such as this is derived in a straightforward way, by log-linearizing equations from modern optimizing macroeconomic models. In fact, the money-demand equation (3) can be obtained directly from a dynamic model in which agents maximize utility of consumption and real balances. Obstfeld and Rogoff (2003) derive such an equation when consumption and real balances enter separably into utility as power functions:

$$\frac{1}{1-\sigma}C_t^{1-\sigma} + \frac{V_t}{1-\varpi}\left(\frac{M_t}{P_t}\right)^{1-\varpi},$$

where V_t is a random shift factor in preferences for real balances. The only difference between the money-demand equation derived from the

first-order condition in the money-in-the-utility function model and the ad hoc money-demand equation is that the log of consumption, c_t, appears as the activity variable in money demand, rather than the log of income, y_t, as in equation (4). Using equation (5), Obstfeld and Rogoff (2003) derive an expression for the exchange rate analogous to equation (7).

Equation (7) is not entirely satisfactory as an "exchange rate equation" because the real exchange rate, q_t, appears as an explanatory variable on the right-hand side of the equation. Equilibrium models with flexible goods prices would elaborate on this equation by relating the equilibrium real exchange rate to underlying economic variables such as productivity and current account balances. Or, the simplest such model, which assumes purchasing power parity, treats q_t as a constant. These models might also relate output differentials to other economic driving variables.

The sticky nominal-price models of Dornbusch (1976) and subsequent authors treat the real exchange rate and perhaps the output differential as endogenous variables whose dynamics are in part determined by the stochastic process for money supplies.

We note also that monetary policy might be endogenous, so that the relative money supplies are set in response to the realizations of macroeconomic variables. We return to discussion of this in much greater detail later.

One might argue that equation (6) holds by definition. That is, equation (4) defines the money-demand errors, $v_t - v_t^*$, and equation (5) defines the deviation from interest parity, ρ_t. The exchange rate equation (6) must hold if we allow a role for these suitably defined unobserved fundamentals. But while we need to acknowledge a role for unobserved fundamentals, this class of models is only interesting if the observed fundamentals do a good job of explaining exchange rates.

As the quotes that we begin this paper with suggest, a standard way of evaluating exchange rate models is to compare their out-of-sample forecasting power to that of the random walk model. There are many variants of the standard model, which might depend on the way the fundamentals are measured, or the set of fundamentals that are included in order to account for the behavior of q_t, $y_t - y_t^*$, $v_t - v_t^*$, and/or ρ_t. Because of the possibility of overfitting or data mining (by a single researcher, or by exchange rate researchers collectively), in-sample fit is considered an unreliable benchmark. Good out-of-sample forecasting power is a

higher hurdle, and has become the standard by which exchange rate models are judged.

Engel and West (2005; hereinafter EW05), however, demonstrate that under some plausible conditions, these models actually have the implication that the exchange rate is "nearly" a random walk. In typical samples, the models actually imply that the change in the exchange rate is not predictable.

The theorem states that as the discount factor, b, goes to 1, the change in the log of the exchange rate between time $t-1$ and t becomes uncorrelated with information in the time $t-1$ information set. The conditions under which that holds are either (a) $\mathbf{a}_1'\mathbf{x}_t$ is integrated of order 1, and $\mathbf{a}_2'\mathbf{x}_t$ is zero, or (b) $\mathbf{a}_2'\mathbf{x}_t$ is I(1).

Note first that this theorem does not require $\mathbf{a}_1'\mathbf{x}_t$ or $\mathbf{a}_2'\mathbf{x}_t$ to be pure random walks. If it did, the theorem would be trivial, since the weighted sum of random walk processes that would appear in the present-value formulas are also random walks.

Second, this theorem does not say that for discount factors less than 1, the log of the exchange rate is exactly a random walk. It says, in essence, that for large values of the discount factor, the log of the exchange rate is approximately a random walk.

To illustrate the theorem, suppose $\mathbf{a}_1'\mathbf{x}_t$ is a scalar x_t, and that $\mathbf{a}_2'\mathbf{x}_t$ is identically zero. As shown in the following, this is a special case of the monetary model. Suppose further that x_t has a unit root but is not a random walk. Assume

$$x_t - x_{t-1} = \phi(x_{t-1} - x_{t-2}) + \varepsilon_t, \varepsilon_t \sim \text{i.i.d.}$$

Then the solution for the change in the exchange rate is given by:

$$s_t - s_{t-1} = \frac{\phi(1-b)}{1-b\phi}(x_{t-1} - x_{t-2}) + \frac{1}{1-b\phi}\varepsilon_t.$$

It is clear from this example that the change in the exchange rate is predictable from the lagged change in the fundamental x_t. But as $b \to 1$, the coefficient on the lagged money supply goes to zero, and the exchange rate approaches a random walk.

The theorem is proved in EW05. Intuitively, suppose first that $\mathbf{a}_2'\mathbf{x}_t$ is identically zero. Consider the discounted sum, $\Sigma_{j=0}^{\infty} b^j \mathbf{a}_1'\mathbf{x}_{t+j}$. Using a Beveridge-Nelson decomposition, we can write $\mathbf{a}_1'\mathbf{x}_t$ as the sum of two components, a pure random walk "permanent" component, and a transitory component, τ_t. We have that $\text{var}_t(\tau_{t+j})$ approaches a constant as j gets large, but the conditional variance of the random walk component

grows in proportion to j. So from the perspective of time t, the permanent component becomes more and more important in accounting for the ex ante variation in $\mathbf{a}_1'\mathbf{x}_{t+j}$ as j gets large. When the discount factor is close to 1, the discounted sum puts a lot of weight on values of $\mathbf{a}_1'\mathbf{x}_{t+j}$ in the future. As $b \to 1$, the discounted sum begins to look more and more like a sum of pure random walk variables.

Now allow non-zero values for both $\mathbf{a}_1'\mathbf{x}_t$ and $\mathbf{a}_2'\mathbf{x}_t$. Because the present value of $\mathbf{a}_1'\mathbf{x}_t$ is multiplied by $(1-b)$, movements in s_t are dominated by the present value of $\mathbf{a}_2'\mathbf{x}_t$ as $b \to 1$. For reasons sketched in the previous paragraph, s_t will behave like a random walk for b near 1 if $\mathbf{a}_2'\mathbf{x}_t$ has a unit root.

If this theorem is applicable to exchange rate models, it suggests that we should not evaluate the models by the criterion of beating a random walk in out-of-sample forecasting power. How close the exchange rate is to a random walk, if it is generated by a present-value model, as in equation (2), depends in practice on how close the discount factor is to 1, and how persistent is the transitory component of the economic fundamentals. EW05 calibrate these for some standard exchange rate models, and show that apparently the models imply near-random walk exchange rate behavior.

For example, consider a simple monetary model in which uncovered interest parity holds ($\rho_t = 0$), purchasing power parity holds ($q_t = 0$), there are no money demand errors ($v_t = v_t^* = 0$), and in which the income elasticity of money demand is unity ($\gamma = 1$). In this case, the exchange rate model simplifies to

$$s_t = \frac{1}{1+\lambda} E_t\left\{\sum_{j=0}^{\infty}\left(\frac{\lambda}{1+\lambda}\right)^j [m_{t+j} - m_{t+j}^* - (y_{t+j} - y_{t+j}^*)]\right\}. \tag{8}$$

A quite conservatively low estimate of λ for quarterly data, from studies of money demand and exchange rates, is $\lambda = 10$, which implies $b \approx 0.90$.[1] In the data for the United States relative to each of the other G7 countries, the highest serial correlation for $\Delta[m_t - m_t^* - (y_t - y_t^*)]$ is 0.41. But the computations of EW05 show that for $b = 0.90$ and a serial correlation of $\Delta[m_t - m_t^* - (y_t - y_t^*)]$ equal to 0.50, that the correlation of Δs_t and Δs_{t-1} is only 0.05, and the correlation of Δs_t with $\Delta[m_{t-1} - m_{t-1}^* - (y_{t-1} - y_{t-1}^*)]$ is 0.06. That is, if the exchange rate were generated from equation (8), it would exhibit near-random-walk behavior. The exchange rate is predictable, but (as EW05 discuss) we would not be likely to reject the random walk in sample sizes that are typically available to open-economy researchers.

We note that the technical conditions for the random walk rule out a stationary process for $\mathbf{a}_2'\mathbf{x}_t$. In the monetary model, this means that the risk premium, ρ_t, if it is present, must not be stationary (see equation [7]). In practice, however, if this term has nearly a unit root, the random walk will nearly follow, as illustrated in the computations in EW05. Nevertheless, in practice, it is a critical question as to whether arguably stationary terms such as risk premia can be exploited to make predictions that beat the random walk. We discuss this further in section 5, when we present results from panel prediction exercises.

We observe that a discount factor close to 1 is helpful to reconcile the observation that the variance in innovations in exchange rates is large relative to the variance of innovations in interest differentials. Using equation (5), and if we associate $\mathbf{a}_2'\mathbf{x}_t$ with the risk premium ($\mathbf{a}_2'\mathbf{x}_t = \rho_t$), we can rewrite equation (1) as:

$$(1 - b)(s_t - \mathbf{a}_1'\mathbf{x}_t) = b(i_t - i_t^*). \tag{9}$$

Typically the volatility of innovations in the fundamentals $\mathbf{a}_1'\mathbf{x}_t$ is small compared to that of exchange rates. But innovations in $s_t - \mathbf{a}_1'\mathbf{x}_t$ have a much higher variance than innovations in the interest differential, $i_t - i_t^*$. In a model such as this, reconciliation of these facts could be accomplished by having the discount factor b close to 1, or by appealing to the claim that there are unobserved components of the fundamentals that have a high variance. It is of course much more satisfying not to have to rely on the volatility of an unobserved variable to account for exchange rate volatility, so models with the discount factor near 1 are appealing on this score.

We turn now to two examples to help elucidate the EW05 theorem.

1.1 Example 1

The Hong Kong dollar (HK\$) per U.S. dollar (US\$) nominal exchange rate is apparently a stationary random variable. It fluctuates between 7.75 and 7.85 HK\$ per US\$. That means that the US\$ per Japanese yen (¥) and HK\$ per ¥ exchange rates are cointegrated (with cointegrating vector [1, –1]). The Engle-Granger representation theorem tells us that if the US\$/¥ and HK\$/¥ exchange rates are cointegrated, then at least one of them is predictable.

We use this example to illustrate the EW05 theorem. Suppose, then, that the US\$/¥ exchange rate s_t^{UJ} is generated by a present-value model,

$$s_t^{UJ} = (1-b)\sum_{j=0}^{\infty} b^j x_t^{UJ},$$

where x_t^{UJ} is an I(1) fundamental. The HK\$/¥ exchange rate, s_t^{HJ}, is determined by an analogous model, with x_t^{HJ} as the I(1) fundamental. If s_t^{UJ} and s_t^{HJ} are cointegrated, then x_t^{UJ} and x_t^{HJ} must also be cointegrated, say $x_t^{UJ} = x_t^{HJ} + z_t$, for some stationary z_t. For simplicity, assume z_t has a mean of zero. Then

$$\begin{aligned} s_t^{HJ} &= (1-b)\sum_{j=0}^{\infty} b^j x_t^{HJ} = (1-b)\sum_{j=0}^{\infty} b^j x_t^{UJ} + (1-b)\sum_{j=0}^{\infty} b^j z_t \\ &= s_t^{UJ} + (1-b)\tilde{z}_t, \end{aligned} \tag{10}$$

where $\tilde{z}_t \equiv \Sigma_{j=0}^{\infty} b^j z_t$ is stationary and has finite variance even in the limit as $b \to 1$. Thus, we can write for the HK\$/US\$ exchange rate, s_t^{HU},

$$s_t^{HU} = s_t^{HJ} - s_t^{UJ} = (1-b)\tilde{z}_t, \tag{11}$$

and as $b \to 1$, $s_t^{HU} \to 0$ (i.e., a constant, equal to zero here since z_t has a zero mean). In other words, the EW05 theorem implies that while s_t^{UJ} and s_t^{HJ} each approach random walks as $b \to 1$, s_t^{HU} approaches a constant.

To get a better sense of the behavior of these exchange rates for $b < 1$, consider the following simple example. Suppose $\Delta x_t^{UJ} = e_t$, $e_t \sim$ i.i.d. (that is, x_t^{UJ} is a random walk), and let z_t defined above also be i.i.d. Then $\tilde{z}_t = z_t$, and $s_t^{HJ} = s_t^{UJ} + (1-b)\tilde{z}_t$, implying

$$\begin{aligned} \Delta s_t^{HJ} &= \Delta s_t^{UJ} + (1-b)\Delta z_t = e_t + (1-b)(z_t - z_{t-1}) \\ &= e_t + (1-b)z_t - (s_{t-1}^{HJ} - s_{t-1}^{UJ}). \end{aligned} \tag{12}$$

So, upon defining the i.i.d. variable $v_t = e_t + (1-b)z_t$, we can write the vector equilibrium correction model (VECM) for s_t^{UJ} and s_t^{HJ} as:

$$\Delta s_t^{UJ} = e_t \tag{13a}$$

$$\Delta s_t^{HJ} = v_t - (s_{t-1}^{HJ} - s_{t-1}^{UJ}). \tag{13b}$$

According to equation (13b), Δs_t^{HJ} is predictable using the lagged HK\$/US\$ exchange rate, $s_{t-1}^{HU} = s_{t-1}^{HJ} - s_{t-1}^{UJ}$. But as $b \to 1$, $s_{t-1}^{HJ} - s_{t-1}^{UJ} \to 0$, and we get the EW05 result that s_t^{HJ} follows a random walk.

1.2 Example 2

Probably the best-cited exchange rate model ever is Dornbusch's (1976) overshooting model. At first glance, it seems as though EW05's theorem

could not apply to Dornbusch's model. The most celebrated aspect of the model—the fact that in response to a permanent money supply shock, the exchange rate overshoots, responding more in the short run than in the long run—implies that exchange rate changes are predictable. When the currency depreciates in response to a domestic monetary expansion, we can predict that it will appreciate toward its long-run equilibrium value.

The EW05 theorem actually is not designed to answer the question of whether the exchange rate in the Dornbusch model theoretically is nearly a random walk. Since equations (3), (4), and (5) hold in the Dornbusch model, then the present value relationship (6) also holds. The EW05 theorem takes the data-generating processes for the fundamentals as given, and asks what the implied exchange rate behavior is for large values of the discount factor. That is subtly different than asking what happens in the model to the behavior of the exchange rate as the discount factor goes to 1, because a change in the discount factor may change the implied data-generating process for the fundamentals. In other words, the EW05 theorem suggests that if the exchange rate is determined by the model (6), with the observed DGPs for the fundamentals, and the discount factor is close to 1, then the exchange rate will be nearly a random walk.

Here we briefly examine the theoretical behavior of the exchange rate in the Dornbusch model when the discount factor is nearly unity, allowing for the fact that the data-generating process for the fundamentals—particularly the real exchange rate—is affected by the discount factor.

We look at a version of the model very close in spirit to Dornbusch's original model. We use equations (3), (4), and (5), and, as in Dornbusch, assume that uncovered interest parity holds exactly, $\rho_t = 0$. As in the overshooting analysis of Dornbusch, we will take output as exogenous. Dornbusch examined the impact of a permanent change in the money supply in a nonstochastic model. In the stochastic setting, this is equivalent to looking at a random-walk process for the money supply. Since output shocks and money demand shocks have identical effects on the exchange rate as money supply shocks (up to the sign of the effect), we will simply assume that the fundamentals follow a random walk:

$$\Delta[m_t - m_t^* - \gamma(y_t - y_t^*) - (v_t - v_t^*)] = u_t, \quad u_t \sim \text{i.i.d.} \tag{14}$$

We need to supplement the model with a price-adjustment equation. The open-economy macro literature of the 1970s and 1980s assumed a backward-looking element to price setting. The log of the domestic price

level for time t, p_t, is preset in time $t-1$, and adjusted to eliminate part of the deviation of p_{t-1} from its long-run equilibrium level. As in Dornbusch, we will assume purchasing power parity (PPP) holds in the long run, so $p_t - p_{t-1}$ eliminates part of the time $t-1$ PPP deviation, $s_{t-1} + p^*_{t-1} - p_{t-1}$. In addition, there is a forward-looking trend term to price adjustment. Obstfeld and Rogoff (1984) emphasize that price-adjustment equations that do not include the forward-looking element lead to counterintuitive dynamics when considering expected future changes in policy, or nonstationary dynamics in the fundamentals. Here, we implement a version of what they call a Mussa rule—the trend term is the expected change in the market-clearing exchange rate. We have:

$$p_t - p_{t-1} = \theta(s_{t-1} + p^*_{t-1} - p_{t-1}) + E_{t-1}s_t + p^*_t - (s_{t-1} + p^*_{t-1}). \tag{15}$$

This pricing rule is symmetric in that the analogous pricing rule for the foreign country yields equation (15) as well. Equation (15) implies that the real exchange rate follows a first-order autoregressive process:

$$E_{t-1}q_t = (1 - \theta)q_{t-1}. \tag{16}$$

Note that the persistence of the real exchange rate is entirely determined by the speed of adjustment of nominal prices. We return to this point below when we discuss the "PPP puzzle." However, the real exchange rate does depend on monetary shocks and on the discount factor, which work through their effect on innovations in the real exchange rate.

If we substitute equations (14) and (16) into the present-value formula (7), we derive:

$$\begin{aligned} s_t - s_{t-1} &= -\frac{\theta}{1 + \lambda\theta} q_{t-1} + \frac{1 + \lambda\theta}{\lambda\theta} u_t \\ &= -\frac{(1 - b)\theta}{1 - b + b\theta} q_{t-1} + \frac{1 - b + b\theta}{b\theta} u_t, \end{aligned} \tag{17}$$

where, recall, the discount factor is $b = \lambda/(1 + \lambda)$. Equation (17) demonstrates the famous overshooting result. In response to a shock, u_t, the exchange rate jumps more than one for one, $(1 + \lambda\theta)/\lambda\theta$. The volatility is greater the stickier are prices (the smaller is θ). But the change in the exchange rate is predictable. When s_t is above its PPP value in period $t-1$ (so q_{t-1} is positive), then $E_{t-1}(s_t - s_{t-1}) < 0$. We can see, however, that the EW05 result holds in this model. As $b \to 1$, $s_t - s_{t-1} \approx u_t$.

As noted earlier, the real exchange rate behavior does depend on the value of b:

$$q_t = (1 - \theta)q_{t-1} + \frac{1 + \lambda\theta}{\lambda\theta} u_t = (1 - \theta)q_{t-1} + \frac{1 - b + b\theta}{b\theta} u_t. \tag{18}$$

Assume that $\lambda = 10$, a conservatively low value if calibrated to quarterly data. Also, assume $\theta = 0.25$, which implies a half life of price adjustment of 2.4 quarters. This speed of price adjustment is in line with the typical calibration of modern sticky-price macroeconomic models. However, it would imply from equation (16) real exchange rate convergence that is much faster than what is typically observed among advanced countries. With these parameters, how well could we predict nominal exchange rate changes using the lagged real exchange rate? The implied *R*-squared from the regression in equation (17) is 0.012, or slightly greater than 1 percent. While the exchange rate change is predictable in theory, in practice it would not be predictable in any reasonably sized sample.

One might expect that if price adjustment were slower (that is, θ lower), that perhaps exchange rate changes would be more predictable. Since the overshooting is greater, the exchange rate has further to adjust to reach its long-run value, and perhaps we can predict that change. In fact, lower values of θ reduce the predictability of exchange rate changes. After a shock to the fundamentals, it is true that the gap between the exchange rate and its long-run value is wider the smaller is θ. The overshooting for a given monetary shock is proportional to $1/\theta$. But, the predictable percentage change in the exchange rate toward its long-run value is smaller in proportion to θ. These two effects precisely offset each other. But the other effect of small values of θ is to make the variance of innovations to the exchange rate larger—the overshooting is larger. So the variance of the unpredictable component of changes in the exchange rate grows relative to the variance of the predictable component as θ gets smaller.

To further develop intuition, we can ask about longer-run changes in the exchange rate. We find:

$$s_{t+k} - s_t = -\frac{[1 - (1 - \theta)^k]}{1 + \lambda\theta} q_t + \sum_{j=1}^{k} \left(\frac{[1 - \theta]^{j-1} + \lambda\theta}{\lambda\theta} \right) u_{t+k+1-j}. \tag{19}$$

There are two effects as the horizon for forecasts gets longer. First, for higher k, the variance of the predictable part of the exchange rate change increases, since $[1 - (1 - \theta)^k]/(1 + \lambda\theta)$ is increasing in k. But, since the exchange rate has a unit root, the variance of the unpredictable part grows without bound as k grows. That can be seen from the second term on the right-hand-side of equation (19), which has a variance that is greater

than k times the innovation variance. In practice, for the parameters considered above ($\lambda = 10$ and $\theta = 0.25$), the maximum R-squared in this regression comes at 6 quarters, and has a value of 0.027. At 16 quarters, the implied R-squared is 0.018.

1.3 *Using Future Values to Forecast: Meese-Rogoff Revisited*

While recent literature has attempted to validate exchange rate models by comparing their out-of-sample forecasting power to that of the random walk, Meese and Rogoff (1983a) did something different. They apparently gave the exchange rate models an advantage relative to the random walk model. Forecasting from the exchange rate model requires forecasting the values of the fundamentals. But Meese and Rogoff evaluated the exchange rate models using the actual realized values of the fundamentals, rather than forecasting them.

The more recent literature has not used the Meese-Rogoff technique, and instead compared monetary exchange rate models to the random walk model using true out-of-sample forecasting power. By this standard, monetary models have not fared well, in general. See, for example, Cheung, Chinn, and Garcia-Pascual (2005), who find that the models generally do not have significantly better forecasting power than the random walk model. In the following, we note that monetary model forecasts based on panel estimation techniques at long horizons (such as in Mark and Sul [2001]) do seem to have greater forecasting power than the random walk. We also examine the forecasting power of models in which monetary policy is endogenized (as in Molodtsova and Papell [2007]). But here we want to reconsider the Meese-Rogoff technique.

As an example (this is one model actually considered by Meese and Rogoff), use equation (4) to solve for the log of the exchange rate, under the assumption of PPP ($q_t = 0$):

$$s_t = m_t - m_t^* - \gamma(y_t - y_t^*) + \lambda(i_t - i_t^*) + u_t, \tag{20}$$

where here the error term u_t is associated with the error terms in money demand, $v_t - v_t^*$.

Rossi (2005) emphasizes that Meese and Rogoff may not have fully accounted for the serial correlation in u_t. She notes that if the exchange rate is cointegrated with the economic fundamentals included in equation (20), then u_t is stationary, but it might be serially correlated. Suppose $u_t = \rho u_{t-1} + \varepsilon_t$, $\varepsilon_t \sim$ i.i.d. Then the model forecast for time $t + 1$ should be

$$m_{t+1} - m^*_{t+1} - \gamma(y_{t+1} - y^*_{t+1}) + \lambda(i_{t+1} - i^*_{t+1}) + \rho u_t.$$

Rossi argues that even with this addendum, the models still might be at a disadvantage relative to the random walk. This is because ρ is plausibly near 1, in which case estimates of ρ tend to be biased downward. Hence imposing a unit root (imposing $\rho = 1$) might result in long-horizon forecasts superior to ones that rely on an estimate of ρ that is far below its true value. Moreover, when ρ is estimated, distributions of test statistics tend to be nonstandard, a result captured by Rossi by modeling ρ as "local to unity."

Suppose we impose $\rho = 1$, then our forecast of s_{t+1} is given by:

$$m_{t+1} - m^*_{t+1} - \gamma(y_{t+1} - y^*_{t+1}) + \lambda(i_{t+1} - i^*_{t+1}) + u_t$$
$$= \Delta[m_{t+1} - m^*_{t+1} - \gamma(y_{t+1} - y^*_{t+1}) + \lambda(i_{t+1} - i^*_{t+1})] + s_t$$

In other words, taking into account the serial correlation of the residual, we should be using the change in the fundamentals to forecast the change in the exchange rate.

Note, however, that Meese and Rogoff (1983b) allow for a grid of possible values for ρ, including $\rho = 1$, and still find that the models do not improve the out-of-sample fit compared to the random walk.

We make a different observation about the Meese and Rogoff standard for evaluating the exchange rate models. There is a sense in which the out-of-sample fit of the model is arbitrary—the model can be rewritten to make the fit arbitrarily good or bad. That is, suppose that x_t is the model for some variable y_t, and we have: $y_t = x_t + \omega_t$. The variance of ω_t, σ^2_ω, serves as a measure of the goodness of fit of the model under the Meese-Rogoff approach. Now, consider rewriting the model as: $y_t = [x_t - (1-a)y_t]/a + \omega_t/a = z_t + \varpi_t$, where a is an arbitrary constant, $z_t \equiv [x_t - (1-a)y_t]/a$, and $\varpi_t = \omega_t/a$. The model is not changed—the second equation is simply an algebraic manipulation of the first. But the variance of the error in the rewritten model—the variance of ϖ_t—is $(1/a)^2$ times the variance in the first model.

For example, using the log approximation to the pure arbitrage-covered interest parity condition, we have $f_t - s_t = i_t - i^*_t$, where f_t is the log of the one-period ahead forward exchange rate. Substitute this expression into equation (20), and rearrange terms to get:

$$s_t = \frac{1}{1+\lambda}[m_t - m^*_t - \gamma(y_t - y^*_t)] + \frac{\lambda}{1+\lambda} f_t + \frac{1}{1+\lambda} u_t. \tag{21}$$

This representation of the model is very similar in spirit to equations (1) or (6): the exchange rate is a weighted average of the (observed) economic fundamentals, and the expected future exchange rate, here measured as f_t. This way of writing the exchange rate equation emphasizes the weight of expectations of the future relative to current fundamentals. But it also gives us an error term with a much lower variance. The variance here is $(1/(1+\lambda))^2$ times the variance of the error term in equation (20). We have been using a value of $\lambda = 10$ for quarterly data in our examples so far, which would imply that the variance of the error term is lower by a factor of 1/121.

Which is the correct way to write the model? Since they are algebraically equivalent, it is hard to argue for one in favor of the other, and indeed that is exactly the problem with the Meese and Rogoff methodology. Moreover, both ways of writing the model have natural economic interpretations. If the model is true with no error, then the right-hand sides of equations (20) and (21) are equal:

$$m_t - m_t^* - \gamma(y_t - y_t^*) + \lambda(i_t - i_t^*)$$
$$= \frac{1}{1-\lambda}[m_t - m_t^* - \gamma(y_t - y_t^*)] + \frac{\lambda}{1+\lambda} f_t. \tag{22}$$

When there is an error term, using the formulation in (20) magnifies the error because it includes $-\lambda s_t$ as an explanatory variable.

To be clear, this critique does not apply to genuine out-of-sample forecast comparisons. At some level, it is obvious that we cannot simply rewrite the model and produce out-of-sample forecasts that have arbitrarily lower variance. If we could, we would not be writing or reading this paper, and instead would be out using this technique to get very rich. While we can rewrite the right-hand side of the model to arbitrarily change the in-sample fit, it follows from equation (22) that our forecast of the right-hand side is the same no matter which way it is written.

Finally, we note that it is often asserted that the forecast using the actual realized values of the explanatory variables must produce better forecasts of the exchange rate than when the right-hand side variables must be forecast. However, this is not in general true if the explanatory fundamental variables are correlated with the unobserved variables. Unless we take this correlation into account, the fit could potentially be worse using the ex post fundamentals.

Our general point, then, is that the Meese-Rogoff procedure of using realized values of the explanatory variables is not invariant to the way

the model is written. Plausible ways of rewriting the model can give much lower mean-squared errors for the model.

Of course, we cannot conclude that if the model is not useful in forecasting exchange rate changes, we have support for the model. Any model can fail to forecast exchange rates. We turn now to alternative means of assessing exchange rate models.

2 Taylor-Rule Models

2.1 Overview

Meese and Rogoff originally suggested the out-of-sample fit criterion as a check on empirical studies that found good in-sample fit. Here, we return to examination of the in-sample fit, but with more attention paid to the market's expectations of future values of the macroeconomic fundamentals.

The monetary models that we have explored so far have been formulated in such a way that the endogeneity of monetary policy has been essentially completely ignored. We have used money supply to capture the monetary fundamental, and have focused on formulations in which nominal interest rates move to equilibrate money supply and money demand. We have not tried to relate movements of the money supply to the macroeconomic variables that policymakers might target.

But modern monetary macroeconomic models formulate the determination of interest rates and monetary equilibrium quite differently. First, they emphasize the endogeneity of monetary policy. Advanced countries have managed to stabilize inflation and apparently establish monetary policy credibility over the past twenty or twenty-five years. If our models of exchange rates are to capture expected future fundamentals, we need to recognize that market forecasts of the future incorporate their assumptions about monetary policy reactions to changes in the macro environment. Second, since the mid-1980s, central banks have used short-term interest rates as their policy instrument, rather than the money supply.

Engel and West (2006) and Mark (2007; hereinafter, EW06 and M07) specify the monetary policy rules for the home and foreign country as interest rate reaction functions for the central bank. Specifically, they assume the home country (in their empirical studies, the home country is Germany, prior to the adoption of the euro) sets the nominal interest rate to target the deviation of expected inflation from the central bank's tar-

get, $E_t\pi_{t+1}$; the output gap, y_t; and, possibly, the deviation of the nominal exchange rate from its purchasing power parity value—that is, the real exchange rate, q_t. The latter term is included to capture the notion that the monetary authorities in some countries tend to raise interest rates when their currency depreciates. For example, Clarida, Galí, and Gertler (1998) find empirical support for this notion in Japan and some other countries. We summarize the monetary policy rule in equation (23):

$$i_t = \gamma_q q_t + \gamma_\pi E_t\pi_{t+1} + \gamma_y y_t + \delta i_{t-1} + u_{mt}. \tag{23}$$

We assume $\gamma_q > 0$, $\gamma_\pi > 1$, $\gamma_y > 0$, and $0 \leq \delta < 1$. Here, u_{mt} represents an error or shift in the monetary policy rule.

The foreign country (taken to be the United States in EW06 and M07) follows a similar policy rule:

$$i_t^* = \gamma_\pi E_t\pi_{t+1}^* + \gamma_y y_t^* + \delta i_{t-1}^* + u_{mt}^*. \tag{24}$$

Here, we assume that the parameters on the inflation deviation and the output gap are the same in the home and foreign country, but assume that the foreign central bank is passive with respect to exchange rate fluctuations.

Using the uncovered interest party (u.i.p.) relationship (including the u.i.p. deviation, ρ_t), $i_t - i_t^* = E_t s_{t+1} - s_t + \rho_t$, we can manipulate these equations to get a forward-looking expression for the real exchange rate:

$$q_t = b\sum_{j=0}^{\infty} b^j E_t z_{t+j}, \tag{25}$$

where

$$b \equiv \frac{1}{1+\gamma_q},$$

$$z_t = -[(\gamma_\pi - 1)(E_t\pi_{t+1} - E_t\pi_{t+1}^*) + \gamma_y(y_t - y_t^*) + \delta(i_{t-1} - i_{t-1}^*) + (u_{mt} - u_{mt}^* - \rho_t)].$$

M07's formulation actually sets $\gamma_q = 0$, because his estimates of the Bundesbank's monetary policy reaction function yielded insignificant estimates for this parameter. With that change, we can still represent the real exchange rate with a present-value expression, but the discount factor, b, is equal to 1. We can do that if z_t has a zero (unconditional) mean—implying q_t has a zero mean, or long-run purchasing power parity holds—and satisfies a mild summability condition. Any ARMA process

satisfies the condition, though we note that fractionally integrated processes do not. In this, but not all, contexts, there will be little difference in empirical results that impose $\gamma_q = 0$ and empirical results that, consistent with the estimates in Clarida et al. (1998), impose small but positive values such as $\gamma_q = 0.10$. Through most of the empirical work in the following, we assume that in quarterly data, γ_q is a small positive number, less than 0.10 but strictly greater than zero.

The equation for the real exchange rate given by equation (25) does not solve a full general equilibrium model in terms of exogenous variables. The fundamentals—inflation and the output gap—are determined by underlying driving variables, such as productivity disturbances and cost-push shocks. We note, nonetheless, that equation (25) has an interesting implication—ceteris paribus, an increase in home relative to foreign inflation leads to a home real appreciation. This prediction stands in contrast to the usual interpretation of the effect of an increase in expected inflation in the monetary models we previously investigated. In those, an increase in expected inflation at home lowers home money demand, leading to a home depreciation. Here, when $\gamma_\pi > 1$, an increase in expected inflation leads the home central bank to raise the home real interest rate, leading to an appreciation.

EW06 and M07 estimate the model as summarized by equation (25). The present-value model requires a measure of expected inflation, output gap, and interest rates for all periods in the future. It also requires an estimate of future expected values of $u_{mt} - u^*_{mt} - \rho_t$. The empirical researcher faces a severe problem in estimating such a model, because the market forms expectations based on many sources of information that are not measurable by the econometrician.

EW06 and M07 handle estimation of the expected present-value sum, (25), in similar ways. First, they ignore $u_{mt} - u^*_{mt} - \rho_t$, treating these as unobservable determinants of the real exchange rate. We now describe EW06's methodology, then note the differences between EW06 and M07. First, EW06 do not include the lagged interest rate in the monetary policy rule, so they have set $\delta = 0$. EW06 then must measure $E_t[(\gamma_\pi - 1)(\pi_{t+j+1} - \pi^*_{t+j+1}) + \gamma_y(y_{t+j} - y^*_{t+j})]$ for all j. EW06 do not estimate the parameters γ_π and γ_y—instead, they base them on estimates of the Taylor rule on post-1979 data in Clarida, Galí, and Gertler (1998). EW06 also use Clarida et al.'s estimate of γ_q, from which the discount factor, b, is calculated. Then, EW06 estimate a VAR in $\pi_{t+1} - \pi^*_{t+1}$, $y_t - y^*_t$, and $i_t - i^*_t$. From the VAR, expected values of $\pi_{t+1} - \pi^*_{t+1}$, and $y_t - y^*_t$ for all periods can be constructed, and then the present value can be calculated.

EW06 then compare the model real exchange rate—the value of the right-hand side of equation (25), where expectations are calculated as just described—to the behavior of the actual deutschemark/dollar real exchange rate. EW06 estimate the model on post-1979 data, using monthly data, 1979:10–1998:12. The correlation of the model real exchange rate and actual real exchange rate is 0.32. That is not extremely high, but it is not too bad, and represents a promising start in this literature. We note, however, that the model real exchange rate estimated by EW06 has a standard deviation about one fifth of that of the actual real exchange rate.

The approach of M07 differs from EW06 in a number of ways. First, M07 estimates the parameters of the Taylor rule, for two periods 1960:II–1979:II, and 1979:III–2003:IV, using quarterly data. Second, as noted previously, Mark does not include the real exchange rate in the interest rate rule for Germany. Third, M07 does include the lagged interest rate in the policy rule. Fourth, M07 compares the behavior of the model real exchange rate to the actual real exchange rate over a longer period, 1976:II–2003:IV.

The correlation of the model and actual real exchange rate in M07 is quite similar to that in EW06, equal to 0.304 (when the output gap is measured using deviations from the HP-filter). However, the model volatility of the real exchange rate is much larger in M07, and more nearly matches that of the real exchange rate. The variance of 1-quarter changes in the real exchange rate from the model is 36.81 percent from the model, compared to 20.06 percent in the data.

2.2 *Purchasing Power Parity Puzzle*

In advanced economies, real and nominal exchange rate changes are highly correlated. A plausible model of real exchange rate behavior must account for this correlation. For an international macroeconomist, a model of nominal exchange rates that cannot be reconciled with real exchange rate behavior is not appealing, and vice versa. Some existing literature (as exemplified by Rogoff [1996]) argues that both sticky and flexible price models fail to replicate some important exchange rate characteristics. As explained in the following, such authors argue that flexible price models have a hard time explaining volatility of real exchange rates, while sticky price models have difficulty explaining persistence of real exchange rates. But recent work by Benigno (2004) and others shows that with suitable modeling of price stickiness and monetary policy, real exchange rate persistence can be plausibly explained as coming from persistence in interest rates.

It is possible to understand the high correlation of real and nominal exchange rates in an environment in which nominal goods prices adjust quickly (no price stickiness), if we assume that monetary authorities stabilize nominal prices. In that case, we should think of nominal exchange rates as being driven by underlying real shocks that drive the real exchange rate. That is, since $s_t = q_t + p_t - p_t^*$, if nominal prices are flexible but p_t and p_t^* are stabilized by monetary policy, then movements in s_t will be highly correlated with those in q_t. However, as Rogoff (1996) emphasized, while flexible-price models can account for the extreme persistence of real exchange rates that we see among advanced countries, they are unable to explain the high volatility.

Alternatively, we might consider models in which p_t and p_t^* have low volatility (that is, low innovation variance) at least in part because of nominal price stickiness. In models with price stickiness, Rogoff notes, we have better explanations of real exchange rate volatility. Specifically, we can appeal to the Dornbusch overshooting model (which we have described in section 1). Monetary shocks cause volatile nominal exchange rate changes—see equation (17). But empirical studies have found the half-life of real exchange rates to be greater than two years. In sticky-price models of the Dornbusch vintage, the sluggishness of real exchange rates is directly tied to the speed of adjustment of nominal prices. Indeed, that can be directly seen from equation (16)—in this version of the model, if the response of real exchange rates to shocks has a half-life of two-plus years, so does the deviation of nominal prices from their flexible-price equilibrium level. Accounting for the slow adjustment of real exchange rates requires unrealistic assumptions about the sluggishness of price adjustment. Rogoff concludes that we cannot easily reconcile real exchange rate behavior with the monetary models of price stickiness, either.

Benigno (2004), however, shows that models with two modifications to the Dornbusch-style model delink real exchange rate persistence and the speed of nominal price adjustment. First, replace the ad hoc price adjustment equation (15) with a Calvo-style price adjustment equation. A fraction of firms each period reset their prices, recognizing that they will not be able to reset prices in every period. They set prices to maximize the value of the firm—the expected present value of current and future dividends. Thus, prices are set in a purely forward-looking manner. Second, monetary policy is endogenous. Central banks set nominal interest rates by a Taylor rule, reacting to the inflation level and the output gap.

We demonstrate the point here in the model from Engel and West (2006), which is directly derived from the working paper version of Galí

and Monacelli (2005). The home country is small relative to the foreign country. That is reflected in the behavior of consumer price inflation in each country. Home-country inflation is a weighted average of inflation of domestically produced goods and home-currency inflation of foreign-produced goods:

$$\pi_t = (1 - \alpha)\pi_{dt} + \alpha\pi_{ft}. \tag{26}$$

The law of one price holds for imported goods in the home country:

$$\pi_{ft} = \Delta s_t + \pi_t^*. \tag{27}$$

This equation embodies the assumption that the home country is small. The term π_t^* represents both the foreign currency rate of inflation of foreign-produced goods and the foreign consumer price inflation. The notion is that the home country is so small relative to the foreign country that while the foreign country does import goods from the home country, the weight of home-country prices is infinitesimally small in the overall foreign consumer price index.

Firms in each country set prices by a Calvo price-setting rule:

$$\pi_{dt} = \beta E_t \pi_{dt+1} + \kappa y_t + u_{ct} \tag{28}$$

$$\pi_t^* = \beta E_t \pi_{t+1}^* + \kappa y_t^* + u_{ct}^*, \tag{29}$$

where y_t and y_t^* represent the output gaps in the home and foreign countries, and u_{ct} and u_{ct}^* are cost-push shocks. The parameter κ captures the speed of price adjustment. In each period, the larger the proportion of firms that are able to adjust their prices, the larger is κ.

In each country, the monetary authority sets the nominal interest rate to react to expected consumer price inflation and the local output gap. In addition, the home country puts some weight on the real exchange rate in its Taylor rule:

$$i_t = \gamma_q q_t + \gamma_\pi E_t \pi_{t+1} + \gamma_y y_t + u_{mt} \tag{30}$$

$$i_t^* = \gamma_\pi E_t \pi_{t+1}^* + \gamma_y y_t^* + u_{mt}^*. \tag{31}$$

We assume parameter values that ensure a stationary solution. Sufficient assumptions are that all parameters are positive, that $\gamma_\pi > 1$, and that $\alpha\gamma_\pi < 1$; u_{mt} and u_{mt}^* are monetary policy errors.

Demand for output from each country depends on consumption levels and the terms of trade. But, in turn, the relative consumption levels are proportional to the real exchange rate through the familiar equilibrium condition that arises under complete markets. The terms of trade

are proportional to the real exchange rate (details are in Galí and Monacelli). We can then write:

$$y_t - y_t^* = \theta q_t + u_{yt} - u_{yt}^*. \tag{32}$$

where $u_{yt} - u_{yt}^*$ represents a relative productivity shock.

Here we briefly summarize how various shocks affect the economic variables in the system:

- A positive monetary policy shock (i.e., an exogenous monetary tightening) causes an appreciation. Inflation and output also decline.
- A positive Phillips curve shock that, given output and expected inflation, transitorily raises inflation and leads to a real appreciation.
- Suppose there is a positive real shock to the IS curve that, given the real exchange rate, raises output. Then in equilibrium, output rises, the real exchange rate falls, and the inflation rate of home-produced goods rises relative to foreign-produced goods.

In order to highlight the comparison of real exchange rate behavior in this model to that of the Dornbusch model, we will set cost push shocks and productivity shocks to zero, leaving only monetary errors, which are assumed to follow an AR(1) process. We define the relative policy error as:

$$u_t = u_{mt} - u_{mt}^*. \tag{33}$$

with

$$u_t = \phi u_{t-1} + \varepsilon_t, \quad 0 \leq \phi \leq 1. \tag{34}$$

Using equations (26)–(34), we get a solution for the real exchange rate:

$$q_t = c_m u_t, \tag{35}$$

where

$$c_m \equiv -(1 - \beta\phi)/d_m < 0$$

$$d_m \equiv (1 - \beta\phi)[\gamma + \eta(1 - \phi)] + \kappa\theta(\gamma_\pi - 1)\phi$$

$$\gamma \equiv \gamma_q + \gamma_y \theta$$

$$\eta \equiv (1 - \alpha\gamma_\pi)/(1 - \alpha).$$

The important thing to note about the solution for the real exchange rate is that it is proportional to the monetary policy error. Real exchange rates have no persistence, except to the extent that there is persistence in

monetary policy errors. In the lingo of the New Keynesian literature, the real exchange rate exhibits no endogenous persistence. Greater price flexibility is represented by a larger value for κ. We see from the expressions for c_m and d_m that as κ grows, the impact effects of shocks falls (c_m falls). So while some price stickiness ($\kappa < \infty$) is required to make the real exchange rate respond to monetary shocks, the degree of price stickiness does not affect the persistence of real exchange rates, but only affects the impact effect of monetary shocks.

In this model, the link between real exchange rate persistence and the speed of adjustment of nominal prices is broken. It might be possible in this simple framework to account for the persistence of real exchange rates, because in practice relative short-term nominal interest rates among advanced countries are very persistent. It is usually not possible to reject a unit root in $i_t - i_t^*$ (at least not statistically). We note here that similar effects on the persistence of the real exchange rate would be generated in the model if shocks to the monetary rule were serially uncorrelated, but the lagged interest rate appeared in the rule with a positive coefficient near 1.

Intuitively, even if there are a large proportion of firms that reset their prices each period, the firms that set their prices must take into account the effects of monetary policy. Suppose there is an expansionary monetary shock. The firms that adjust their price cannot raise prices fully to the level they would attain if all prices were set flexibly, because their prices might then be quite high relative to firms that have not adjusted prices this period. As argued for the closed-economy model of West (1988), the combination of persistence in monetary shocks and nonsynchronized price setting can stretch out the price adjustment process, even if a large fraction of firms adjust prices at any given time.

The simple model presented here is not realistic enough to describe the most important moments in the macroeconomic data. For example, the output gap and inflation inherit the same persistence as the real exchange rate. Benigno (2004) shows how various types of asymmetries can contribute to even more sluggish real exchange rate adjustment—asymmetries in the price-setting rules, in monetary policy, and so forth.

The other aspect of the PPP puzzle is real exchange rate volatility. Again, we do not attempt to calibrate the previous simple model to see if it is able to generate realistic volatility for plausible values of the parameters. However, Benigno (2004) and Benigno and Benigno (2006) do undertake such exercises in somewhat richer models, and find success in matching both the persistence and volatility of real exchange rates.

We do not mean to suggest that the PPP puzzle is solved. For example, estimated open-economy DSGE models frequently do not produce parameter estimates that will fully account for the persistence and volatility of real exchange rates (see Jung [2007], for example). But we do believe that the Taylor-rule models provide a fruitful direction for future research. In the open-economy empirical literature, much effort has been put into reconciling the speed of adjustment of real exchange rates to the speed of adjustment of prices. Various authors have suggested that the half-life of real exchange rates might be overstated because of estimation bias (see, for example, Murray and Papell, 2002, who emphasize the imprecision in estimates of the speed of reversion of real exchange rates. However, Choi, Mark, and Sul [2006] find an unbiased point estimate of the half life of three years, with a 95 percent confidence interval between 2.3 to 4.2 years.) Others have suggested that aggregation bias can account for sluggish real exchange rate adjustment. That is, individual goods prices might adjust more quickly than the real exchange rate because the CPI-based real exchange rate aggregates individual prices, and aggregates adjust more slowly (Imbs, Mumtaz, Ravn, and Rey 2005). Others have attempted to reconcile the behavior of prices and real exchange rates by appealing to nonlinearities (e.g., Taylor, Peel, and Sarno 2001), or to some sort of slow adjustment of nominal exchange rates, perhaps due to transaction costs in foreign exchange markets (Cheung and Lai 2004). What is intriguing about the open-economy models based on Taylor rules and Calvo price adjustment is that there may be no need to reconcile sluggish real exchange rate adjustment with the frequency of price setting.

3 Granger-Causality Tests and Variance Bounds

We have emphasized that in models encapsulated by equations (1) and (2), the market's expectations of future fundamentals are the key to understanding exchange rate movements. But we are then faced with the dilemma of measuring expectations. We can follow the tack taken by Engel and West (2006) and Mark (2007), and use the forecasts from a statistical model for the fundamentals as measures of the market's expectations. But the market surely uses more information than is contained in a simple VAR forecast, so this method mismeasures expectations.

In this section, we examine two alternative approaches to evaluating present-value models of exchange rates. First, we examine whether exchange rates can help forecast future (observed) fundamentals. Second,

we ask whether the present value of observed fundamentals is sufficiently volatile to account for observed exchange rate volatility. In section 4, we look at alternative methods of measuring expectations.

3.1 *Forecasting Fundamentals*

If the exchange rate reacts to news about future fundamentals, then perhaps the exchange rate is useful in forecasting those fundamentals. Econometric evaluation of this observation was pioneered by Campbell and Shiller (1987). But the Campbell-Shiller analysis is not directly applicable to exchange rate models. This is because future fundamentals typically include some variables that are unobservable, even ex post. As explained in EW05, the present-value models do not necessarily imply that the exchange rate will Granger-cause observed fundamentals. But if the unobserved fundamentals are not the primary drivers of exchange rates, then perhaps movements in exchange rates are useful to forecast macroeconomic variables such as relative money supplies, outputs, prices, or interest rates. That is, while the presence of unobserved fundamentals breaks the tight restriction tested by Campbell and Shiller, it is possible that the exchange rate might Granger-cause the standard observed fundamentals. For this reason, we examine whether exchange rates Granger-cause the fundamentals from the money and Taylor-rule models.

Following the example illustrated by EW05, set $\mathbf{a}_2'\mathbf{x}_t$ from equation (1) equal to zero for simplicity, and write $\mathbf{a}_1'\mathbf{x}_t = x_{ot} + x_{ut}$, where x_{ot} stands for the fundamentals whose ex post values are observed by the econometrician, and x_{ut} signifies the unobserved fundamentals. We can rewrite equation (2) as:

$$s_t = (1 - b)E_t\left[\sum_{j=0}^{\infty} b^j(x_{ot+j} + x_{ut+j})\right]. \tag{36}$$

In section 5, we present evidence from panel cointegration tests that s_t is cointegrated with the observed fundamentals from the monetary and Taylor-rule models. In other words, the unobserved fundamentals, x_{ut}, are found to be stationary. Equation (36) suggests that the exchange rate might contain news about future observed fundamentals, x_{ot}, though it also is affected by news about future unobserved fundamentals. We ask, then, whether the exchange rate can help forecast the observed fundamentals, in the sense that it Granger-causes them.

We write the relationship between the observed fundamentals and the log of the exchange rate in error-correction form:

$$s_t - s_{t-1} = \alpha_1 + \beta_1(x_{ot-1} - s_{ot-1}) + \sum_{i=1}^{I} \gamma_{1i}(s_{t-i} - s_{t-1-i}) + \sum_{j=1}^{J} \delta_{1j}(x_{ot-j} - x_{ot-1-j}) \tag{37a}$$

$$x_{ot} - x_{ot-1} = \alpha_2 + \beta_2(x_{ot-1} - s_{ot-1}) + \sum_{i=1}^{I} \gamma_{2i}(s_{t-i} - s_{t-1-i}) + \sum_{j=1}^{J} \delta_{2j}(x_{ot-j} - x_{ot-1-j}) \tag{37b}$$

The null that the exchange rate does not Granger-cause the fundamentals is represented by the restriction $\beta_2 = \gamma_{21} = \gamma_{22} = \{\ldots\} = \gamma_{2J} = 0$. If we reject this null, it means we accept the hypothesis that the exchange rate is helpful in forecasting future values of x_{ot}. Conversely, in order to accept the hypothesis that the observed fundamentals Granger-cause the exchange rate, we must reject the null $\beta_1 = \delta_{11} = \delta_{12} = \{\ldots\} = \delta_{1J} = 0$. We set the lag length J to 4.

In table 6.1, we report results of these tests. We try three permutations of the observed fundamentals from the monetary model: $\Delta(m_t - m_t^*) - \Delta(y_t - y_t^*)$, which follows from the monetary model when the income elasticity of money demand is unity; and $\Delta(m_t - m_t^*)$ and $\Delta(y_t - y_t^*)$ separately.

We also examine whether the exchange rate can help forecast fundamental variables implied by the Taylor-rule models. Rewrite the Taylor-rule model as:

$$i_t - i_t^* = \gamma[s_t - (p_t - p_t^*)] + u_t. \tag{38}$$

Here, we have simply rolled into the disturbance term all the variables that might be targeted by the central bank other than the real exchange rate. Then using interest parity (and including the deviation from interest parity into the general disturbance term), we can rewrite equation (38) in one of two ways:

$$s_t = \frac{\gamma}{1+\gamma}(p_t - p_t^*) - \frac{1}{1+\gamma}u_t + \frac{1}{1+\gamma}E_t s_{t+1} \tag{39}$$

$$s_t = \gamma(p_t - p_t^*) + \gamma(i_t - i_t^*) - u_t + (1-\gamma)E_t s_{t+1}. \tag{40}$$

These models suggest using $\Delta(p_t - p_t^*)$ and $\Delta(i_t - i_t^*)$ as measures of the fundamentals.

We perform the tests on quarterly data for the United States plus

Table 6.1
Granger-Causality Tests

	Austria	Australia	Belgium	Canada	Denmark	Finland	France	Germany	Greece
				H0: Δs fails to cause Δx_o					
$i - i^*$	N/A		**		N/A	NA	**	**	
$m - m^*$		*	**						
$m - m^* - (y - y^*)$		***	***			*			
$p - p^*$		*	***		**			*	*
$(y - y)$		**	*		**				
				H0: Δx_o fails to cause Δs					
$i - i^*$	N/A		**	**	N/A	N/A		**	**
$m - m^*$	*		***				**	**	**
$m - m^* - (y - y^*)$									
$p - p^*$	*		**			*	*	**	*
$(y - y)$		**				*			

	Gr. Britain	Italy	Japan	Korea	Netherlands	Norway	Spain	Sweden	Switzerland
				H0: Δs fails to cause Δx_o					
$i - i^*$	**		***	N/A	N/A	N/A			**
$m - m^*$									
$m - m^* - (y - y^*)$				**					
$p - p^*$		***	*	***					
$(y - y)$				***	*			***	
				H0: Δx_o fails to cause Δs					
$i - i^*$	***	*	*	N/A	N/A	N/A	**	**	*
$m - m^*$	**				***		**	**	***
$m - m^* - (y - y^*)$	**	***		**	**			**	**
$p - p^*$	*	**		***				*	**
$(y - y)$	**	***			**	**			**

Notes: The data are quarterly: 1973:1–2005:4. In variables i^*, m^*, y^*, and p^*, the "*" denotes data from one of the eighteen countries listed in the headers to the columns, while variables with no star refer to the United States. Variable definitions: i and i^* are Treasury bill rates; m and m^* are the sum of money plus quasi-money, with exceptions noted in the text; y and y^* are industrial production; p and p^* are the CPI; s is the end of quarter nominal exchange rate versus the U.S. dollar. All variables except i and i^* are measured in logarithms. In euro-area countries, post-1998 data on exchange rates were glued to the earlier data, with suitable normalization. For further details see section 3. We estimated bivariate error-correction models in $(\Delta s_t, \Delta x_{ot})'$, where Δs_t is the change in the (log) of the nominal exchange rate and x_{ot} is one of the five variables listed in the rows of each of the two panels. Each equation in the bivariate model included a constant, a lag of $x_{ot} - s_t$, and four lags of Δs_t and Δx_{ot}. The table reports the results of F-tests on all four lags of Δs_t are zero in the equation for Δx_{ot}; the bottom panel reports parallel tests for lags of Δx_{ot} in the Δs_t equation. Rejections are indicated at the 10% (*), 5% (**) and 1% (***) levels.

eighteen other OECD countries.[2] We use an update of the Mark and Sul (2001) data, beginning in 1973:I, and extended through the end of 2005. Nominal exchange rates are from the IFS CD-ROM (line code AE), end-of-quarter observations. Exchange rates for the euro countries after 1998 were glued onto the euro-dollar rate with the appropriate normalization. We used quarterly industrial production indices for all countries as a proxy for national income because quarterly GDP is unavailable for several countries in the sample. The IMF's International Financial Statistics (IFS line code 66) provides our measure of industrial production. Our measure of money is from the IFS and is the sum of money (line code 34) plus quasi-money (line code 35) for all countries with the following exceptions for Great Britain, Norway, and Sweden, due to availability. Money is M0 from the IFS for Great Britain, M2 from the OECD's Main Economic Indicators for Norway, and M3 also from the OECD for Sweden. The IFS continues to report currency and demand deposits and other deposits for euro-zone countries after the introduction of the euro. Price levels are measured using the CPI from the IFS (line code 64).

Mark and Sul (2001) did not use data on interest rates. We use line 60c (Treasury bill rate) from IFS as a source for short-term interest rate. However, for some countries, interest rate data were not available for much of our time span from this source. For these countries, we report NA in table 6.1 for causality tests involving interest rates. However, for Japan, we were able to get the data on interest rate differential from Todd Clark, who has kindly provided the dataset used in Clark and West (2006).

Here, following Engel and West (2005), we conduct the tests country by country (rather than using panel estimation), using the Akaike criterion to determine lag length.

The tables show that the exchange rate does have modest power in forecasting future fundamentals. This finding is certainly not uniform across all fundamentals and across all countries. The forecasting power seems more uniformly good across countries for the fundamentals based on Taylor rules: relative prices and relative interest rates.

The second panel in each table shows results from the reverse causality test: whether the fundamentals forecast the change in the exchange rate. Here we find a surprisingly large number of rejections of the null—that the fundamentals do not Granger-cause changes in exchange rates. This evidence is in conflict with much of the literature, which has found the models have little power to forecast exchange rate changes. In section 5, we discuss circumstances under which the models might be use-

ful in forecasting exchange rate changes (that is, we discuss possible cases in which the conditions of the EW05 theorem do not hold). We explore the models' ability to forecast exchange rates in that section using the more rigorous criterion of out-of-sample forecasting power.

3.2 *Volatility*

Let us define:

$$h_{It} = (1 - b)\sum_{j=0}^{\infty} b^j E(x_{t+j} \mid \Omega_{It}), \quad 0 < b < 1$$

where h_{It} is the expected present discounted value of x_{t+j}. Expectations are taken with respect to the information set Ω_{It}, which is the market's information at time t. We think of this present value as being a component of the determinants of the exchange rate. Specifically, h_{It} is associated with the present value of the macroeconomic fundamentals that drive exchange rates that can be observed by the econometrician.

As we have stated earlier informally, the econometrician may measure this present value with error, because the econometrician's information set at any time t, Ω_{Jt}, is contained in the market's information set. The econometrician can calculate:

$$h_{Jt} = (1 - b)\sum_{j=0}^{\infty} b^j E(x_{t+j} \mid \Omega_{Jt}).$$

Engel and West (2004) demonstrate that if x_t is an I(1) random variable, as the discount factor, b, gets close to 1, $\mathrm{var}(h_{Jt} - h_{Jt-1}) \approx \mathrm{var}(h_{It} - h_{It-1})$. This theorem does *not* say that the econometrician's measure, h_{Jt}, gets close to h_{It} as $b \to 1$. But the result can be used in the following way: suppose we observe only some subset of the fundamentals that drive the exchange rate, which we call x_t. By calculating h_{Jt} (where we estimate $E[x_{t+j} \mid \Omega_{Jt}]$ from a statistical model, in this case a 4th-order autoregression in Δx_t), we can derive a measure of $\mathrm{var}(h_{Jt} - h_{Jt-1}) \approx \mathrm{var}(h_{It} - h_{It-1})$. We can then ask how $\mathrm{var}(h_{It} - h_{It-1})$ compares to $\mathrm{var}(s_t - s_{t-1})$; that is, how much of the variance of the change in the exchange rate can be accounted for by the variance of the change in h_{It}.

To be sure, this calculation cannot tell us whether the present value h_{It} is a good model of the exchange rate. It merely seeks to answer the question of whether the observed fundamentals are volatile enough to account for the volatility in the exchange rate (as measured by the variance of the change in the exchange rate). This question is of some interest, be-

cause exchange rates are known to be very volatile, so it is natural to ask whether there is excess volatility. Here, we do not mean to ask whether the variance of changes in the log of the exchange rate have a greater variance than can be explained by the present-value model, because we cannot observe all of the fundamentals that belong in the present value. Instead, in essence, we ask whether the observed fundamentals can account for enough of exchange rate volatility that they are plausible candidates for explaining highly volatile nominal exchange rates.

We first consider the monetary fundamentals: $x_t = m_t - m_t^* - (y_t - y_t^*)$. Table 6.2 reports $\mathrm{var}(h_{jt} - h_{jt-1})/\mathrm{var}(s_t - s_{t-1})$. We report results for values of the discount factor b equal to 0.90, 0.95, and 0.99.

The table shows that the monetary fundamental can generally account for a fairly high fraction of exchange rate variance. The case of Canada is unusual because the measure of the variance ratio is greater than 1, so that if the model is correct, the unobserved fundamentals would have to be negatively correlated with the monetary fundamentals. Excluding the case of Canada, the variance ratio is on average equal to 0.48 when $b = 0.95$.

We can write the Taylor-rule model of equation (25) as a model for nominal exchange rates:

$$s_t = \Phi_t + \Gamma_t, \tag{41}$$

where

$$\Phi_t = (1 - b)\sum_{j=0}^{\infty} b^j E_t(p_{t+j} - p_{t+j}^*), \tag{42}$$

$$\Gamma_t = \sum_{j=0}^{\infty} b^j E_t(z_{t+j} - \pi_{t+j+1} + \pi_{t+j+1}^*),$$

and b and z_t are defined as in equation (25). (Equation [41] expresses s_t in the form of equation [2], with Φ_t and Γ_t corresponding to equation [2]'s present values in $\mathbf{a}_1'\mathbf{x}_t$ and $\mathbf{a}_2'\mathbf{x}_t$, respectively.)

In the second column of table 6.2, we treat Γ_t as an unobserved fundamental and assume that z_t contains a unit root. We are then relying only on Φ_t to account for the volatility of the exchange rate, and not Γ_t. The Φ_t component accounts for less of the variance than we found with the monetary model. Again, take for example the case of $b = 0.95$. The present value of the fundamentals can account for only the fraction 0.19 (averaged across all countries) of the exchange rate variance.

We leave to future work the examination of the model when Γ_t is

Table 6.2
Variance Ratios

		fundamental	
	b	$m-y-(m^*-y^*)$	$p-p^*$
Australia	0.9	0.349	0.152
	0.95	0.364	0.228
	0.99	0.378	0.356
Austria	0.9	0.327	0.079
	0.95	0.347	0.079
	0.99	0.363	0.157
Belgium	0.9	0.449	0.065
	0.95	0.515	0.108
	0.99	0.589	0.194
Canada	0.9	6.981	0.147
	0.95	6.832	0.169
	0.99	6.705	0.191
Denmark	0.9	0.528	0.045
	0.95	0.554	0.053
	0.99	0.58	0.063
Finland	0.9	0.538	0.139
	0.95	0.594	0.24
	0.99	0.651	0.458
France	0.9	0.326	0.063
	0.95	0.353	0.111
	0.99	0.381	0.219
Germany	0.9	0.412	0.053
	0.95	0.481	0.087
	0.99	0.561	0.156
Greece	0.9	0.605	0.306
	0.95	0.62	0.543
	0.99	0.637	1.12
Gt. Britain	0.9	0.458	0.185
	0.95	0.54	0.283
	0.99	0.642	0.456
Italy	0.9	0.527	0.196
	0.95	0.597	0.343
	0.99	0.682	0.667
Japan	0.9	0.481	0.034
	0.95	0.551	0.044
	0.99	0.627	0.059

(*continued*)

Table 6.2
Continued

		fundamental	
	b	$m-y-(m^*-y^*)$	$p-p^*$
Korea	0.9	0.549	0.151
	0.95	0.611	0.205
	0.99	0.678	0.28
Netherlands	0.9	0.252	0.154
	0.95	0.272	0.156
	0.99	0.293	0.158
Norway	0.9	0.302	0.112
	0.95	0.283	0.165
	0.99	0.268	0.251
Spain	0.9	0.559	0.206
	0.95	0.627	0.366
	0.99	0.701	0.737
Sweden	0.9	0.433	0.098
	0.95	0.44	0.134
	0.99	0.45	0.186
Switzerland	0.9	0.431	0.047
	0.95	0.442	0.078
	0.99	0.453	0.14

Notes: Data are described in notes to table 6.1. Let $x_t = m_t - m_t^* - (y_t - y_t^*)$ or $x_t = p_t - p_t^*$, let b be one of the three discount factors listed in the table, and let $h_{jt} = (1-b)\Sigma_{j=0}^{\infty} b^j E(x_{t+j} \mid \Omega_{jt})$, where the expectation is computed using a fourth-order autoregression in Δx_t. The table reports estimates of $\mathrm{var}(h_{jt} - h_{jt-1})/\mathrm{var}(s_t - s_t)$. See their "Volatility" subsection in section 3 of the paper for more details.

included. We note that it is unlikely that the condition that z_t have a unit root is satisfied in our data for OECD countries, but if z_t is highly persistent—so that its largest root is near unity—our method of calculating the implied volatility should work in practice.

4 Using Surveys to Measure Expectations

4.1 Announcement Studies

Present-value models as characterized by equation (2) have the implication that news about current and future fundamentals influence exchange rates. We could examine the plausibility of models by testing

whether exchange rates react to news about future fundamentals in the way predicted by the models.

When government agencies announce data measurements—GDP growth, unemployment rates, inflation, and so on—economic agents learn about the current value of fundamentals, and also revise their expectations of future fundamentals. "News" is the information contained in the announcement—the difference between the announced level of the economic variable and the market's expectation. The problem confronting researchers, as we have noted, is measuring expectations. We can proceed (as in, for example, the studies of the Taylor-rule model by Engel and West [2006] and Mark [2007], discussed previously) by inferring expectations from the estimates of a VAR or other statistical model. However, it is likely that the market uses many other sources of information to form expectations that are contained in typical statistical analyses, so the researcher may mismeasure expectations.

An alternative means of measuring expectations is from survey data. In particular, several recent studies have examined the response of exchange rates to announcements of economic news, and use a measure of the market's expectation of that announcement culled from surveys of market participants. In particular, these studies have typically used the surveys conducted by Money Market Services (MMS). Since 1977, MMS has conducted a survey each Friday of some 40 money managers at commercial and investment banks, recording forecasts of all indicators to be released (in the United States) in the subsequent week. The news contained in the release is measured as the difference between the announced value and the median forecast of that value by the MMS survey.

At least four recent papers have made use of high-frequency data (Faust et al. 2003, 2007; Andersen et al. 2003, and Clarida and Waldman 2007). These studies examine the response of the exchange rate from shortly before to shortly after the announcement (for example, in a ten-minute window). Many earlier studies performed similar studies (such as Engel and Frankel 1984, and Hardouvelis 1988), but using exchange rate changes measured at less-high frequencies (for example, from the open of the New York market to the close of the New York market on the day of the announcement).

The recent studies uniformly find responses of dollar exchange rates in line with the predictions of the Taylor-rule model. Specifically, news of activity variables that suggest an expanding economy in the United States, or announcements of higher inflation (greater than expected),

lead to an appreciation of the dollar at the time of the announcement. To see this, write out equation (25), omitting the lagged interest rates and unobserved variables for convenience:

$$q_t = -\sum_{j=0}^{\infty}\left(\frac{1}{1+\gamma_q}\right)^{j+1} E_t[(\gamma_\pi - 1)(\pi_{t+j+1} - \pi^*_{t+j+1}) + \gamma_y(y_{t+j} + y^*_{t+j})]. \tag{44}$$

Recall the restrictions on the coefficients: $\gamma_q > 0$, $\gamma_\pi > 1$, $\gamma_y > 0$.

Denote the exchange rate immediately before the announcement as q_{t-}, and immediately after, q_{t+}, and use a similar notation to capture the change in expectations. Then we have:

$$q_{t+} - q_{t-} = -\sum_{j=0}^{\infty}\left(\frac{1}{1+\gamma_q}\right)^{j+1} (E_{t+} - E_{t-})[(\gamma_\pi - 1)(\pi_{t+j+1} - \pi^*_{t+j+1}) + \gamma_y(y_{t+j} - y^*_{t+j})]. \tag{45}$$

From this expression, we see that the dollar should appreciate (q_t falls) if U.S. inflation is announced to be greater than expected, or if news revises expectations upward about the future U.S. output gap.

Andersen et al. (2003) examined the response of the value of the dollar in terms of Swiss francs, (synthetic) euros, British pounds, Japanese yen, and German marks from January 3, 1992, through December 30, 1998. They look at the reaction to a wide range of macroeconomic announcements. The findings go strongly in the direction predicted by the Taylor-rule model. For example, for all five exchange rates, the dollar appreciates (on average) to positive news about U.S. GDP, nonfarm payrolls, retail sales, industrial production, durable goods orders, construction spending, factory orders, and several other indicators of economic activity. If we interpret this news as meaning that agents revise their opinion about the current and future output gaps in the United States upward, the Taylor-rule model indeed posits an appreciation of the dollar. The mechanism is that the Fed is expected to increase interest rates, which makes U.S. assets more attractive, inducing a dollar appreciation to equilibrate the asset market. Strikingly, it is also the case that the dollar appreciates against all five currencies with news about PPI or CPI inflation: when inflation in the United States is greater than expected, the dollar appreciates.

Faust et al. (2007) look at the response of the German mark/euro and the British pound to macroeconomic announcements between January 9, 1987, and December 18, 2002, again for several macro variables. Their findings are very similar to those of Andersen et al.: For both currencies,

the dollar appreciates in response to positive news about U.S. economic activity. When there is a positive surprise to GDP, housing starts, nonfarm payrolls, retail sales, or a negative surprise to initial unemployment claims or unemployment, the dollar appreciates. Moreover, the dollar appreciates to positive inflation surprises to the PPI for both currencies, and to the CPI for the dollar/pound rate.

Clarida and Waldman (2007) examine the reaction of the dollar value of the Japanese yen, Canadian dollar, Norwegian krone, Swedish krona, Swiss franc, euro, British pound, Australian dollar, and New Zealand dollar from July 2001 to December 2005. This study differs from the previous two in that it uses data on macroeconomic announcements from non-U.S. countries. In particular, it looks at the reaction to news about inflation from each of these countries, as measured by the surprise component of the announcement of consumer price inflation. Expectations are measured from surveys conducted by Bloomberg news service.

The specification examines how the local currency responds to news about local inflation. In a regression that pools across all currencies, Clarida and Waldman find that an announcement of higher inflation (greater than expected) leads to a stronger currency. The findings carry over to country-by-country regressions, though they are a bit weaker in the two countries (the United States and Japan) that do not explicitly target inflation. Indeed, in England and Norway, there is evidence that the effect changed as the countries changed monetary policy to target inflation.

Faust et al. (2003) examine a slightly different situation: they look at the effects of surprise changes in the Fed's target for its policy instrument—the Fed funds rate—using high-frequency data. They look at the response of the dollar value of the mark-euro and the pound to 62 announcements from March 1994 to October 2001. The expected Fed funds rate is measured from Fed funds futures data. As the models would predict, there is a strong and significant effect on the currencies, where a surprise increase in the Fed funds rate appreciates the dollar.

Engel and Frankel (1984) looked at the effects of announcements of the money supply in the United States in the early 1980s on the value of the dollar. That study related the change in the exchange rate to the difference between the announced value of the money supply and the expected value of the announced money supply as calculated from the MMS survey of forecasters. The flavor of the finding is similar to the recent studies—an unexpectedly high U.S. money supply led to an appreciation of the dollar. Why? The markets apparently believed that the

central bank was likely to react to this announcement by contracting the money supply. Indeed, Engel and Frankel also found that short-term interest rates reacted positively to the money surprise. In the early 1980s, the Fed (supposedly) had a money-supply target, so the reaction of the real exchange rate reflected the credibility of Fed policy. Likewise, the exchange rate reaction to news about inflation in the Clarida and Waldman paper reflects the credibility of inflation targeting central banks.

Hardouvelis (1988) looks at the reaction of interest rates and exchange rates to announcements of a number of economic variables in the October 1979–August 1984 period. His findings are consistent with both Engel-Frankel and Clarida-Waldman. That is, he finds that the dollar consistently and significantly appreciates in response to positive surprises in the money supply, as in Engel-Frankel, and as we would expect if the Fed were credibly targeting the money stock. But the reaction of the dollar to announcements of CPI and PPI inflation is mixed. Across seven exchange rates, the sign of the response varies and is never statistically significant. This finding contrasts with the more recent literature, which finds that positive home inflation surprises almost uniformly lead to a stronger home currency. But we note that the period of Hardouvelis's study is one in which the Fed probably had not yet established the credibility it had in the 1990s and later for fighting inflation.

4.2 Long-Run Expectations Based on Surveys

In order to directly evaluate the model of the real exchange rate given by equation (44), we need a measure of expected inflation and output gap for all periods into the future. Rather than measuring expectations of macroeconomic fundamentals far into the future by using VARs, we consider survey measures of expectations. Consensus Forecasts publishes a monthly volume that surveys economic forecasters, and constructs an average forecast for a large number of economic and financial variables. Twice a year (in April and October), for some of these variables, they canvass forecasters for their forecast of what the variable will take in the current year, in each of the next five years, and then for an average for years 6 through 10 from the current year.

We use these surveys to construct the present-value relationship in equation (44). We use the surveys from April 1997 to October 2006 for the United States, Japan, Germany, France, the United Kingdom, Canada, Italy, the Netherlands, Norway, Spain, Sweden, and Switzerland.[3] We use forecasts of inflation and output.

There are three important limitations to this data:

1. The survey asks for expectations of output but the model requires expectations of the output gap. The expected relative output gap, $E_t(y_{t+j} - y^*_{t+j})$, corresponds to expected relative output for all j when the full employment level of output in the home and foreign country remain in constant proportion.

2. The surveys record expectations only 10 years into the future, while the sum in equation (44) requires expectations infinitely far into the future. To handle the truncation, for inflation we assume that the expectations recorded for years 6–10 for each country will hold into the indefinite future. However, it is unrealistic to assume that expected output growth differentials will continue forever. The survey provides output growth expectations for years 6–10 for each country. We assume (at each survey date) that the market expects the output growth for years 11 onward for each country to equal the average output growth expected for all countries.

3. There is not much data. We have observations only twice a year, for 10 years. We estimate the model in first differences, which removes country-fixed effects. So we have at most 19 observations for each country. For a few countries, there are some missing data points and we have fewer observations.

For each date, we construct a present value of forecasts of current and future inflation, and current and future output, using a semiannual discount factor of 0.9. We choose this value based on estimates of γ_q in the monetary policy rule in Germany by Clarida, Galí, and Gertler (1998). We construct the present values for output and inflation for the United States relative to each of the other countries.

We estimate the regression for real exchange rates in first differences:

$$q_t - q_{t-1} = a + b_t(PDV_{yt} - PDV_{yt-1}) + b_2(PDV_{\pi t} - PDV_{\pi t-1}) + b_3(q_{t-1} - q_{t-2}) + u_t. \tag{46}$$

Here, PDV_{yt} refers to the present discounted values of relative output expectations and $PDV_{\pi t}$ to the present discounted value of relative inflation expectations. We began with a specification implied by the model, with $b_3 = 0$, but found serial correlation in the residuals. We reestimated, including the lagged dependent variable, and we report results from both regressions in the following.[4]

We estimate this equation by OLS, but acknowledge two potential

problems: first, it is likely that there is still serial correlation in u_t. Suppose u_t is the first difference in the expected discounted sum of the errors in the monetary policy rule above, $(u_{mt} - u^*_{mt} - \rho_t)/(1 + \gamma_q)$. With the real exchange rate sampled every six months, expected changes in the monetary rule for one year or more in the future will result in serially correlated values for u_t. This problem is partly attenuated by the inclusion of the lagged change in the real exchange rate in the regression. We also have ignored the possibility of correlation between the error term and the explanatory variables in the regression, leading to coefficient estimates that might be inconsistent.

We estimate equation (44) for each of the 11 U.S. real exchange rates. The coefficients in the regressions are only occasionally significant and so we do not report them. However, this is not surprising, given the paucity of data. We note that in the country-by-country regressions, the coefficient on the relative-output present values have the correct sign in 10 of 11 cases, and the coefficient on the relative-inflation present value is correctly signed in 9 of 11 cases.

To increase efficiency, we estimated the model as a panel. We note that estimating in differences eliminates country-specific fixed effects to levels of the real exchange rate. We estimated the panel model in three versions: with no intercept, with a common intercept, and with country-specific intercepts. In the two specifications with intercept terms, the estimated intercepts were small and statistically insignificant in all cases. The estimated coefficients on the discounted sums were essentially unaffected by the inclusion of these intercept terms, and the standard errors of the estimates increased a very small amount when intercepts were included. The results are presented for the no-intercept case in tables 6.3a and 6.3b. Note especially the coefficient on the discounted sum of current and future expected inflation—it is negative. As we have emphasized, when central bank policy to target inflation is credible, an increase in home relative to foreign inflation leads to a real home appreciation.

There are obvious limitations to this empirical exercise. Aside from the data limitations as described, researchers must use survey data on expectations with a modicum of caution. We really do not have any idea whether the forecasts from the surveys do correspond to the forecasts of market participants, and there is also the possibility that the professional forecasters are not reporting their true expectations accurately. (For example, they may "talk their book.") On the other hand, the limitations of using forecasts from VARs are also well recognized: the mar-

Table 6.3a
Regression Results, $b_3 = 0$

	Coefficients	Standard Error
Y discounted	–0.6106	0.2052
π discounted	–3.8850	2.3606

Table 6.3b
Regression Results, $b_3 \neq 0$

	Coefficients	Standard Error
Y discounted	–0.8802	0.2115
π discounted	–4.3405	2.2901
q lagged	0.2522	0.0678

kets use more information than is contained in the VAR, there may be shifts in the data-generating process not captured by the VAR, and so on. It is interesting and reassuring that the model fits well using this data on long-run expectations.

5 Out-of-Sample Forecasts Based on Panel Estimates

5.1 Motivation

Mark and Sul (2001), Rapach and Wohar (2002), and Groen (2005) have in fact found that panel error-correction models (ECM) based on the simple monetary model using the fundamental $m_t - m_t^* - (y_t - y_t^*)$, or the closely related purchasing power parity model, in which the fundamental is given by $p_t - p_t^*$, do have the power to forecast exchange rates out of sample. The forecasting power is particularly strong at long horizons.

In this section, we replicate Mark and Sul (2001) with a longer sample, confirming the finding that the random walk can be beaten in a panel study. We do not have a precise economic or econometric story as to why panel estimates do so much better than country-by-country estimates. We conjecture that the efficiency of panel estimation is key.

In any event, we know from the argument in section 1 that beating the random walk requires a stationary, unobserved component to the fundamentals. Consistent with the monetary model in equation (7), we label this component ρ_t, and our discussion interprets this as a risk pre-

mium. On the one hand, the EW05 result shows that acceptable formulations of the monetary model imply that the exchange rate should look very much like a random walk. The calculations in EW05 find that the exchange rate would be nearly unforecastable if it is entirely determined by the observable fundamentals, given the observed data-generating processes for these fundamentals. But it is also possible that, in an ECM framework, the monetary model could be used to forecast changes in exchange rates, especially at longer horizons, when the unobserved ρ_t is present.

Consider this example, which is very similar to the type of model that Campbell (2001) uses to account for long-horizon predictability in stock prices:[5]

Suppose the exchange rate is determined by:

$$s_t = (1 - b)x_t + b\rho_t + bE_t s_{t+1}.$$

Assume the fundamental x_t is observed, and for simplicity assume it follows a random-walk process:

$$x_t - x_{t-1} = \varepsilon_t, \quad \varepsilon_t \sim \text{i.i.d.}, \sigma_\varepsilon^2 \equiv \text{var}(\varepsilon_t).$$

We could assume that the first difference is serially correlated, as in example 1 in section 1. When b is close to 1, the observed fundamental imparts near-random-walk behavior to the exchange rate. Our point here is made more transparent by simply assuming the observed fundamental is a random walk.

Now assume the deviation from uncovered interest parity, ρ_t, which is the unobserved fundamental, and follows a stationary first-order autoregressive process:

$$\rho_t = \alpha\rho_{t-1} + u_t, \quad u_t \sim \text{i.i.d.}, \sigma_u^2 \equiv \text{var}(u_t).$$

The forward-looking no-bubbles solution for the exchange rate is:

$$s_t = x_t + \frac{b}{1 - b\alpha}\rho_t. \tag{47}$$

Assume that the econometrician does not have an independent observation of the risk premium. However, if the model is correct, then the econometrician can observe the "error," $z_t \equiv x_t - s_t$, which is an indirect observation on ρ_t, since

$$z_t = x_t - s_t = \frac{-b}{1 - b\alpha}\rho_t. \tag{48}$$

Now, suppose we are interested in forecasting the k-period change in the log of the exchange rate, using an error-correction model. We have:

$$s_{t+k} - s_t = x_{t+k} - x_t + \frac{b}{1 - b\alpha}(\rho_{t+k} - \rho_t) \tag{49}$$

$$= \sum_{j=1}^{k} \varepsilon_{t+j} + \frac{b}{1 - b\alpha}(\alpha^k - 1)\rho_t + \frac{b}{1 - b\alpha}\sum_{j=1}^{k} \alpha^{k-j} u_{t+j}$$

$$= \sum_{j=1}^{k} \varepsilon_{t+j} + (1 - \alpha^k) z_t + \frac{b}{1 - b\alpha}\sum_{j=1}^{k} \alpha^{k-j} u_{t+j}.$$

How good are forecasts of the change in the exchange rate at the k-horizon? We can use the theoretical R-squared from the above error-correction model, assuming the parameters are known, to gauge this. We calculate this R-squared for the k-period ECM as:

$$R_k^2 = \frac{(\alpha^k - 1)^2 \operatorname{var}(x_t)}{\operatorname{var}(s_{t+k} - s_t)} \tag{50}$$

$$= \frac{(1 - \alpha^k)^2 \sigma_u^2}{(1 - \alpha^k)^2 \sigma_u^2 + (1 - \alpha^{2k})\sigma_u^2 + [k(1 - \alpha^2)(1 - b\alpha)^2 \sigma_\varepsilon^2 / b^2]}.$$

This R-squared may take on a humped-shape pattern in k. For the shortest horizons, it may be small, but may initially increase as k increases. However, as k goes to infinity, the R-squared goes to zero.

Calibrating this model is difficult because we do not have a good measure of the variance of innovations to ρ_t, or a measure of its serial correlation. Suppose that the ratio of standard deviations of the innovations of the observed fundamental to ρ_t is 3: that is, $\sigma_\varepsilon / \sigma_u = 3$. Suppose further that the serial correlation of ρ_t in quarterly data is 0.95, and set the discount factor as $b = 0.90$. Then we find that $R_1^2 = .02$, meaning little short-run predictability. But initially, as the horizon increases, the predictability from the ECM rises. At 16 quarters, we have $R_{16}^2 = 0.21$, and the maximum R-squared occurs at 44 quarters, with a value of $R_{44}^2 = 0.28$. It seems possible that we would not detect any short-run forecasting power for the model in this case, but that at longer horizons the forecasting power would become more apparent. As we shall see, this intuition is not well borne out when we estimate forecasting equations country by country; by and large, there is not much more predictability at long ($k = 16$ quarters) than at short ($k = 1$ quarter) horizons. But when we estimate a panel, and forecasts rely in part on mean reversion in panel estimates of time effects, we find distinctly more predictability at

16 than at 1 quarter horizons. Perhaps this occurs because all of the exchange rates are dollar exchange rates, and there is a common element to the risk premium. That is, $\rho_{it} = \bar{\rho}_t + \varsigma_{it}$ for exchange rate i, so that the panel estimation is picking up the common component of the risk premium, $\bar{\rho}_t$.

5.2 *Forecasting Methodology*

Here we update the previous studies. It has been claimed that the forecasting power of models is sensitive to the sample (for example, see Faust, Rogers, and Wright (2003), so it is useful to reexamine previous studies using more current data.

Our econometric analysis centers on panel estimation of the short-horizon predictive regression,

$$s_{it+k} - s_{it} = \beta_k z_{it} + \varepsilon_{it+k}, \tag{51}$$

where

$$z_{it} \equiv x_{it} - s_{it}, \text{ and } \varepsilon_{it} \equiv \zeta_i + \theta_t + u_{it}.$$

Here, i indexes the country and t is the time period. We give the regression error ε_{it} an unobserved components interpretation where ζ_i is an individual-specific effect, θ_t is a time-specific effect that allow us to account for a limited amount of t cross-sectional dependence, and u_{it} is the residual idiosyncratic error. The error-correction term is $z_{it} \equiv x_{it} - s_{it}$, representing the deviation of the exchange rate from the fundamentals for country i. Monetary fundamentals are defined as in example 1,

$$x_{it} \equiv m_{0t} - m_{it} - \gamma(y_{0t} - y_{it}), \tag{53}$$

with $\gamma = 1$ and where the United States serves as the base country (denoted by '0'). Alternatively, PPP fundamentals are defined as:

$$x_{it} = p_{0t} - p_{it}. \tag{34}$$

We also consider fundamentals based on the Taylor rule. As in Molodtsova and Papell (2007), we develop an error-correction formulation for the Taylor rule model by noting that under uncovered interest parity, $E_t s_{it+1} = i_{0t} - i_{it} + s_{it}$. We then replace $i_{0t} - i_{it}$ with the components of the (relative) Taylor rules. So here we use $i_{0t} - i_{it} = 1.5(\pi_{0t} - \pi_{it}) + 0.1(y^g_{0t} - y^g_{it}) + 0.1(s_{it} + p_{it} - p_{0t})$, a Taylor rule that we posit (rather than estimating the coefficients). Here, y^g is the output gap, computed with an HP filter. The parameter on the inflation gap is fairly standard (and coincides

with the estimated Taylor rules in M07). We have used a slightly smaller coefficient on output gap than is typically used, on the grounds that many of the countries in our sample have announced pure inflation-targeting regimes. We have used a parameter of 0.10 on the real exchange rate as in our Taylor-rule models discussed previously. Thus we have:

$$x_{it} = 1.5(\pi_{0t} - \pi_{it}) + 0.1(y^g_{0t} - y^g_{it}) + 0.1(s_{it} + p_{it} - p_{0t}) + s_{it}. \tag{55}$$

The Taylor-rule model then gives us $E_t s_{it+1} - s_{it} = x_{it} - s_{it}$. Following Molodtsova and Papell (2007), we use the forecasting model $E_t s_{it+k} - s_{it} = \beta_k(x_{it} - s_{it})$. Molodtsova and Papell find that the Taylor-rule model is able to forecast well at short horizons using univariate methods, while finding little support for the monetary or PPP models at any horizon.

The predictive regression is most appealing when z is I(0), or equivalently when x and s are cointegrated. Since we are not estimating the cointegration vector, we can test for cointegration with standard unit-root tests. We use Sul's (2006) RMA test, in which the null hypothesis being tested is that z_{it} is I(1) for all i (details are in the notes to table 6.4). The results reported in that table indicate that we reject the null of no cointegration for PPP and Taylor-rule fundamentals but not monetary fundamentals. In the interest of simplicity, however, and in light of the pos-

Table 6.4
Panel Unit Root Tests

	Monetary		PPP		Taylor rule	
	t-stat	p-value	t-stat	p-value	t-stat	p-value
HO: unit root in common factor	–1.155	>0.10	–1.897	<.05	–2.141	<.05
H0: unit roots in error correction terms in all 18 countries	na		–4.713	<.001	–6.356	<.001

Notes: Data are described in notes to table 6.1. Let an "i" subscript denoting country, with the United States defined to be country 0 ($i = 0$). In country i ($i \neq 0$), let x_{it} be one of three measures of fundamentals: $x_{it} = m_{0t} - m_{it} - (y_{0t} - y_{it})$ (*monetary* fundamentals); $x_{it} = p_{0t} - p_{it}$ (*PPP* fundamentals); $x_{it} = 1.5(\pi_{0t} - \pi_{it}) + 0.1(y^g_{0t} - y^g_{it}) + 0.1(s_{it} + p_{it} - p_{0t}) + s_{it}$ (Taylor-rule fundamentals, where y^g is the HP-filtered output gap). For any one of the measures of fundamentals defined in the previous note, let $z_{it} = x_{it} - s_{it}$ denote the corresponding error-correction term. The table report's Sul's (2006) panel unit root test. The procedure first tests whether one can reject a unit root in a common factor. If that null is rejected, one performs a panel unit root test of the null that z_{it} is I(1) for all eighteen countries. The regressions include a time trend and individual country-fixed effects.

sible low power of unit-root tests, we forecast using z as previously defined for monetary as well as PPP and Taylor-rule fundamentals.

We generate out-of-sample forecasts both at a short-horizon ($k = 1$) and at a long-horizon ($k = 16$). We begin by estimating the predictive regression by LSDV on observations available through 1982.4. The $k = 1$ regression is then used to forecast the 1-quarter-ahead exchange rate return in 1983.1 and the $k = 16$ regression to forecast the 16-quarter-ahead exchange rate return through 1986.4. We then update the estimation sample by one period by adding the observation for 1983.1 and repeat the procedure. This recursive updating scheme gives us 92 $k = 1$ forecasts and 77 overlapping $k = 16$ forecasts. We compare the panel regression forecasts against those implied by the random-walk model. We do this for both for the monetary, PPP, and Taylor rule fundamentals. In the Taylor-rule forecasts, we compute the output gap with an HP filter, using only the data in the estimation sample.

Because the time effect enters contemporaneously in the predictive regression, the k-period-ahead forecast requires a forecast of the time effect. We use the recursive mean of the time effect as this forecast and form the exchange-rate prediction according to

$$s_{it+k} - s_{it} = \hat{\zeta}_i + \left(\frac{1}{t}\sum_{j=1}^{t}\hat{\theta}_i\right) + \hat{\beta} z_{it}. \tag{56}$$

We measure relative forecast accuracy with Theil's U-statistic—the ratio of the root-mean-square prediction error (RMSPE) from two competing models: $U < 1$ if forecasts from the predictive regression are more accurate than the random walk. We report t-statistics from the Clark and West (2006, 2007) procedure for testing for equal predictive ability of two nested models.

The data are the same as used in section 3, except that we control for seasonality by filtering the money and output series by applying a one-sided moving average of the current observation and 3-lagged values.

5.3 *Results*

The prediction results for the United States are displayed in tables 6.5a, 6.5b, and 6.5c. We compare the predictive power of the models against two versions of the random walk: with drift and without drift. Engel and Hamilton (1990) and Diebold, Gardeazabal, and Yilmaz (1994) make the case for using the random-walk-with-drift null. Engel and Hamilton present evidence that one can reject the null that the drift in the exchange

rate of the dollar (against the yen, mark, and pound) is constant in their sample. They then note, "The driftless random walk is just a special case of this hypothesis. Imposing a particular value for the drift (in this case, zero) may of course improve the fit over selected subsamples" (p. 701). That is, using a zero-drift random walk as the null imposes a restriction that does not come from theory, but instead from peeking at the data. However, Meese and Rogoff (1983a) did use the driftless random walk as the null, and indeed, in our sample the driftless random walk outperforms the random walk with drift in forecasting exercises. So we report the performance of the model against both null hypotheses.

At the 1-quarter horizon, the monetary fundamentals (table 6.5a) produce a lower RMPSE than the random walk with drift for thirteen of the eighteen currencies, but they outforecast the random walk with no drift for seven of the eighteen exchange rates. In only four cases does the monetary model significantly (at the 10 percent level, using the 1-sided t-test based on the Clark-West statistic) outforecast the random walk with no drift at the short horizon, but the model performs significantly better than the random walk with drift for nine exchange rates.

The PPP fundamentals (table 6.5b) also generally produce lower RMPSEs than the random walk with drift (in fifteen out of eighteen cases), but not so for the driftless random walk (a lower RMPSE in only eight of eighteen cases), at the 1-quarter horizon. In general, the performance of the PPP model and the random walk are very similar at the 1-quarter horizon, with U-statistics very near unity. The PPP model produces significantly lower RMPSEs than the random walk with drift for six exchange rates, but it outperforms the driftless random walk significantly in only three cases at the 1-quarter horizon.

The Taylor-rule model's performance (table 6.5c) at the 1-quarter horizon is similar to that of the PPP model. The RMPSE of forecasts from the Taylor-rule model are very close to those from the random walk, both with and without drift. The Taylor-rule model actually produces a lower RMPSE than the random walk with drift for fourteen of the eighteen currencies, but does so for only two currencies for the driftless random walk. Of the fourteen cases in which the Taylor-rule model yields lower RMPSEs than the random walk with drift, only three are significantly lower.

The forecasting performance of the models relative to the random walk at the 1-quarter horizon is not very good, which is exactly what the EW05 theorem predicts. However, the picture changes when we look at the 16-quarter-ahead forecasts.

Table 6.5a
Panel Data Error-Correction Forecast Evaluation Using Monetary Fundamentals

	No Drift				Drift			
	u(1)	cw	u(16)	cw	u(1)	cw	u(16)	cw
Australia	1.027	0.696	1.307	0.674	1.018	0.224	1.029	0.904
Austria	**0.985**	**1.772**	**0.749**	**2.001**	**0.981**	**1.808**	**0.720**	**3.390**
Belgium	1.007	0.691	**0.788**	**1.910**	1.000	**1.284**	**0.670**	**3.165**
Canada	1.178	–0.959	1.093	**1.601**	1.160	–1.362	**0.901**	**1.662**
Denmark	1.010	0.102	1.042	0.797	**0.999**	1.002	**0.822**	**2.014**
Finland	1.003	0.684	**0.974**	0.886	**0.991**	**1.292**	**0.783**	**1.465**
France	1.005	0.232	**0.825**	**1.605**	**0.991**	**1.286**	**0.609**	**4.314**
Germany	**0.988**	**1.517**	**0.623**	**2.448**	**0.983**	**1.771**	**0.576**	**3.583**
Greece	1.011	**2.346**	1.360	**1.578**	1.031	–0.973	1.263	–0.138
Gt. Britain	1.004	0.334	1.010	0.747	**0.990**	**1.330**	**0.676**	**2.731**
Italy	1.013	0.672	1.174	0.621	**0.986**	**1.523**	**0.739**	**1.441**
Japan	**0.998**	0.760	**0.855**	**1.429**	**0.999**	1.190	**0.849**	**3.257**
Korea	**0.997**	0.874	**0.865**	0.596	**0.987**	**1.428**	**0.718**	0.763
Netherlands	**0.991**	**1.291**	**0.751**	**2.143**	**0.985**	**1.536**	**0.684**	**3.771**
Norway	1.005	0.270	**0.897**	**1.607**	**0.996**	1.173	**0.734**	**3.635**
Spain	1.037	0.472	1.356	0.760	1.008	0.951	**0.826**	**1.725**
Sweden	**0.992**	1.094	**0.658**	**2.227**	**0.980**	**2.308**	**0.510**	**5.549**
Switzerland	**0.985**	**1.949**	**0.675**	**2.766**	**0.978**	**2.033**	**0.658**	**4.634**
Average	1.013		0.945		1.004		0.765	

Notes: Data are described in notes to table 6.1. For quarterly horizons $k = 1$ and $k = 16$, the panel regression $s_{it+k} - s_{it} = \beta_k z_{it} + \varepsilon_{it+k}$ was estimated. See notes to table 6.4 for definition of the error-correction term z_{it} for monetary (tables 6.5a and 6.6a), PPP(tables 6.5b and 6.6b), and Taylor-rule (tables 6.5c and 6.6c) fundamentals. The regressions include a time effect and fixed effects. The sample period for the dependent variable in the initial regression sample is 1973:2–1982:4 ($k = 1$) or 1976:1–1982:4 ($k = 16$). The regression estimates (including estimates of time and fixed effects) were used as described in the text to predict the 1-quarter exchange rate change through 1983:1 or the 16-quarter change through 1986:4. An observation is then added to the end of the sample and the procedure is repeated. This recursive updating scheme yields 92 $k = 1$ predictions and prediction errors and 77 overlapping $k = 16$ predictions and prediction errors. The "u(1)" and "u(16)" columns report Theil *U*-statistics, presenting the ratio of the root mean-squared prediction error (RMSPE) from the procedure described above relative to the RMSPE from a random walk or random walk with drift. Values less than one are in **bold,** and indicate that the panel error-correction model had a lower RMSPE than did the random walk. The drift parameter in the random walk with drift was computed recursively, using the samples described above. The "cw" columns represent Clark and West (2006, 2007) statistics for the null of equal predictive accuracy. Values greater than 1.282 are in **bold** and indicate rejection at the 10 percent level using a one-sided test.

Table 6.5b
Panel Data Error-Correction Forecast Evaluation Using PPP Fundamentals

	No Drift				Drift			
	u(1)	cw	u(16)	cw	u(1)	cw	u(16)	cw
Australia	**0.996**	1.014	**0.868**	**1.667**	**0.987**	**1.505**	**0.683**	**5.596**
Austria	**0.995**	0.980	**0.590**	**2.824**	**0.991**	1.194	**0.568**	**3.359**
Belgium	1.003	0.407	**0.707**	**2.586**	**0.996**	1.035	**0.602**	**4.031**
Canada	1.013	0.384	**0.861**	**1.827**	**0.997**	**1.407**	**0.709**	**4.276**
Denmark	1.006	–0.007	**0.886**	**1.412**	**0.995**	1.068	**0.699**	**3.124**
Finland	1.001	0.520	**0.774**	**1.400**	**0.989**	**1.395**	**0.622**	**2.640**
France	1.010	–0.226	**0.895**	**1.448**	**0.996**	0.963	**0.660**	**3.710**
Germany	**0.999**	0.928	**0.652**	**2.489**	**0.994**	1.105	**0.604**	**2.800**
Greece	**0.997**	**2.335**	1.220	**1.534**	1.018	–0.337	1.133	0.373
Gt. Britain	1.006	0.430	1.063	0.580	**0.991**	1.065	**0.711**	**1.575**
Italy	1.024	0.460	1.312	0.580	**0.997**	0.779	**0.826**	1.106
Japan	**0.991**	**1.410**	**0.736**	**2.880**	**0.992**	1.151	**0.732**	**4.559**
Korea	**0.985**	**1.941**	**0.808**	0.795	**0.976**	**2.632**	**0.670**	1.151
Netherlands	1.016	0.535	**0.755**	**2.508**	1.011	0.642	**0.687**	**2.953**
Norway	**0.999**	0.537	**0.896**	1.080	**0.990**	1.248	**0.734**	**2.056**
Spain	1.035	0.093	1.400	0.487	1.006	0.498	**0.852**	**1.356**
Sweden	1.002	0.425	**0.736**	**1.545**	**0.989**	**1.652**	**0.571**	**3.135**
Switzerland	**0.995**	1.232	**0.576**	**2.272**	**0.987**	**1.404**	**0.562**	**2.159**
Average	1.004		0.874		0.995		0.701	

Notes: See notes to table 6.5a.

The monetary model (table 6.5a) produces lower RMPSEs than the random walk with drift for all but two currencies (Greece and Australia) at the 16-quarter horizon, though for only eleven of the eighteen currencies when the null is the random walk without drift. The average *U*-statistic across all of the eighteen currencies is quite low, 0.765. The monetary model does very well in forecasting some currencies. The monetary model's forecast is significantly better than the random walk with drift for fifteen currencies, and significantly better than the random walk without drift for eleven currencies. In the latter case, the average *U*-statistic is 0.945.

It is interesting to examine the models in which the random walk with no drift performs better than the random walk with drift. The British pound is a case in point—the ratio of the RMPSE for the random walk with no drift to the random walk with drift is 1.541 (=.9405/.6102). The monetary model outforecasts the random walk with drift, but not the

Table 6.5c
Panel Data Forecast Evaluation Using Taylor-Rule Fundamentals

	No Drift				Drift			
	u(1)	cw	u(16)	cw	u(1)	cw	u(16)	cw
Australia	1.005	0.615	1.047	0.541	**0.996**	**1.410**	**0.824**	1.060
Austria	1.002	0.238	**0.891**	1.197	**0.998**	0.859	**0.857**	**1.312**
Belgium	1.006	–0.875	1.051	–0.189	**0.999**	0.516	**0.894**	**1.957**
Canada	1.010	0.059	1.120	0.186	**0.995**	1.234	**0.923**	0.740
Denmark	1.010	–0.962	1.235	–0.726	**0.999**	0.769	**0.974**	0.435
Finland	1.008	–0.634	1.137	–0.254	**0.997**	1.272	**0.914**	0.845
France	1.012	–0.617	1.290	–0.599	**0.999**	0.615	**0.952**	0.509
Germany	1.004	0.072	**0.937**	0.907	**0.999**	0.667	**0.868**	1.224
Greece	**0.989**	**2.461**	1.175	**1.631**	1.010	–1.740	1.091	–0.073
Gt. Britain	1.013	–0.212	1.345	–0.308	**0.998**	0.658	**0.900**	0.529
Italy	1.027	0.313	1.549	–0.068	1.000	0.152	**0.975**	0.204
Japan	**0.997**	0.833	**0.856**	1.161	**0.998**	0.703	**0.850**	0.733
Korea	1.005	0.603	1.176	0.308	**0.995**	**1.474**	**0.976**	0.146
Netherlands	1.008	–0.275	**0.976**	0.987	1.003	–0.150	**0.888**	**2.012**
Norway	1.007	–0.650	1.103	–0.026	**0.998**	0.835	**0.903**	0.677
Spain	1.030	0.039	1.682	–0.402	1.001	–0.155	1.024	0.043
Sweden	1.009	–0.145	1.159	0.085	**0.996**	**1.387**	**0.898**	0.662
Switzerland	1.006	0.392	**0.832**	1.112	**0.998**	0.935	**0.811**	0.679
Average	1.008		1.142		0.999		0.918	

Notes: See notes to table 6.5a.

driftless random walk. It would have taken a courageous forecaster to choose the random walk with no drift as the forecasting model in 1983. The pound depreciated from a price of $2.58 in June 1973 (and $2.42 in October 1980) to $1.09 in February 1985. It surely must have seemed like there was downward drift in the dollar/pound ratio at that point, but indeed, the pound eventually appreciated to $1.87 in May 1988, then fluctuated and reached a value of $1.96 in December of 2006. It is clear why the random-walk model with no drift does better than the random walk with drift if the drift were estimated over the 1973–1985 period, but that clarity comes perhaps only with hindsight. We have performed recursive *t*-tests currency by currency for the null of no drift in the random walk using 16-quarter changes.[6] Roughly speaking, at any given date between 1982 and 2005, we would reject the null of no drift for about half the exchange rates.

The PPP model (table 6.5b) produces a lower RMPSE than the random

walk with drift for all currencies except the Greek drachma. In this case, the improvement is significant in fifteen of the cases, and the average *U*-statistic is 0.701. Compared to the random walk with no drift, the PPP model produces a lower RMPSE in fourteen of the eighteen cases, and it is significantly better for thirteen of the currencies. The average *U*-statistic is 0.874, and 0.770 across the fourteen currencies for which the RMPSE was lower than the random walk. The PPP model is able to deliver large improvements in forecast accuracy relative to the random walk at longer horizons.

The results are mixed at the 16-quarter horizon for the Taylor rule model (table 6.5c). The model produces RMPSEs that are lower than the random walk with drift for sixteen currencies, but only three are significantly lower. The random walk with no drift outperforms the Taylor-rule model in terms of RMPSE for thirteen of the eighteen currencies.

Table 6.6 produces the analogous statistics to table 6.5, but only

Table 6.6a
Single-Equation Error-Correction Forecast Evaluation Using Monetary Fundamentals

	No Drift				Drift			
	u(1)	cw(1)	u(16)	cw(16)	u(1)	cw(1)	u(16)	cw(16)
Australia	1.042	–0.307	1.777	0.657	1.032	–1.395	1.398	**1.383**
Austria	1.010	–0.090	1.290	0.654	1.004	0.773	1.203	**1.823**
Belgium	1.006	0.347	1.895	**1.931**	**0.999**	0.916	1.612	**2.764**
Canada	1.030	1.017	1.509	0.455	1.014	0.882	1.243	–0.082
Denmark	1.011	–0.301	1.431	**1.328**	**0.999**	0.524	1.129	**2.368**
Finland	1.014	–0.571	1.548	0.869	1.003	0.106	1.244	1.161
France	1.003	0.785	1.083	**2.326**	**0.989**	**1.785**	**0.799**	**4.107**
Germany	1.000	0.804	1.136	**2.071**	**0.995**	**1.295**	1.052	**2.989**
Greece	**0.990**	**2.533**	1.669	**1.482**	1.010	–0.973	1.549	0.091
Gt. Britain	1.010	0.131	2.029	0.748	**0.995**	0.960	1.358	**2.064**
Italy	1.016	0.906	1.287	1.054	**0.990**	**1.402**	**0.810**	**1.526**
Japan	1.008	–0.344	1.282	0.305	1.008	0.120	1.274	**1.771**
Korea	1.017	1.057	1.187	0.489	1.007	**1.377**	**0.986**	0.335
Netherlands	1.000	0.773	**0.949**	**1.530**	**0.995**	**1.293**	**0.864**	**2.605**
Norway	1.006	0.427	1.536	**1.604**	**0.997**	0.856	1.257	**3.085**
Spain	1.040	0.387	1.974	**1.373**	1.011	0.065	1.202	**2.170**
Sweden	**0.998**	0.819	1.641	**1.876**	**0.985**	**1.806**	1.272	**4.005**
Switzerland	**0.994**	1.206	1.056	1.202	**0.987**	**1.675**	1.030	**2.760**
Average	1.011		1.460		1.001		1.182	

Notes: See notes to table 6.5a.

Table 6.6b
Single-Equation Error-Correction Forecast Evaluation Using PPP Fundamentals

	No Drift				Drift			
	u(1)	cw(1)	u(16)	cw(16)	u(1)	cw(1)	u(16)	cw(16)
Australia	1.021	–0.376	1.270	0.487	1.011	–0.018	**0.999**	**2.391**
Austria	1.002	0.293	**0.968**	**2.512**	**0.996**	0.874	**0.903**	**3.163**
Belgium	1.006	0.307	1.565	**2.431**	**0.999**	0.630	1.331	**3.424**
Canada	1.011	–0.439	1.162	0.623	**0.996**	1.007	**0.957**	**2.183**
Denmark	1.009	–0.110	1.251	**1.857**	**0.997**	0.748	**0.987**	**3.267**
Finland	1.009	–0.306	**0.923**	1.278	**0.998**	0.702	**0.742**	**2.042**
France	1.013	0.122	1.496	**2.099**	**0.999**	0.595	1.104	**3.973**
Germany	1.002	0.561	1.361	**2.634**	**0.997**	0.843	1.261	**2.861**
Greece	**0.992**	**2.787**	1.824	1.036	1.013	0.041	1.693	0.641
Gt. Britain	1.013	0.213	1.361	0.693	**0.999**	0.568	**0.911**	1.267
Italy	1.036	–0.137	1.708	0.782	1.009	–0.745	1.075	1.078
Japan	1.003	0.396	1.044	0.964	1.003	–0.100	1.037	**2.401**
Korea	**0.979**	**2.462**	**0.883**	1.006	**0.970**	**2.942**	**0.733**	**1.427**
Netherlands	1.012	–0.038	1.194	**2.660**	1.006	0.058	1.087	**3.004**
Norway	1.007	0.183	1.389	**1.424**	**0.998**	0.672	1.137	**2.510**
Spain	1.037	0.011	1.735	0.798	1.008	–1.139	1.056	**1.393**
Sweden	1.020	0.539	1.987	**2.158**	1.007	0.669	1.540	**4.047**
Switzerland	1.000	1.095	**0.782**	**2.356**	**0.992**	1.154	**0.762**	**2.376**
Average	1.010		1.326		1.000		1.073	

Notes: See notes to table 6.5a.

when the models are estimated country by country rather than by panel. It is clear from the table that the panel estimates improve the forecasts relative to the single-equation estimation for the monetary and PPP models in terms of RMPSE. However, the Taylor-rule model performs about as well with single-equation estimation as with the panel. Perhaps this occurs because the restrictions imposed by the panel—that the model is identical across all countries—are far too strong in this case, because monetary rules differ too much across the set of eighteen countries.

There is one interesting aspect of table 6.6 that requires comment. There are several cases (for example, see the 16-quarter-ahead forecast for Sweden for the PPP model, reported in table 6.6b), in which the Clark-West t-statistic indicates that the model significantly outperforms the random walk (both with drift and without drift), even though the

Table 6.6c
Single-Equation Error-Correction Forecast Evaluation Using Taylor-rule Fundamentals

	No Drift				Drift			
	u(1)	cw(1)	u(16)	cw(16)	u(1)	cw(1)	u(16)	cw(16)
Australia	1.010	0.700	1.004	0.742	1.001	0.328	0.790	1.269
Austria	1.007	0.119	**0.679**	**2.671**	1.001	0.412	**0.633**	**3.320**
Belgium	1.013	–0.147	**0.845**	**2.578**	1.005	–0.117	**0.719**	**3.992**
Canada	1.015	–0.035	1.145	0.275	1.000	0.278	**0.944**	0.729
Denmark	1.011	–0.738	1.200	0.223	1.000	0.235	**0.947**	1.119
Finland	1.025	–0.694	1.032	0.594	1.014	–0.555	**0.829**	1.274
France	1.024	–0.033	1.268	1.004	1.011	–0.160	**0.936**	**2.291**
Germany	1.013	–0.037	**0.779**	**2.398**	1.008	–0.034	**0.721**	**2.517**
Greece	**0.984**	**2.633**	1.187	**1.658**	1.004	0.681	1.103	–0.138
Gt. Britain	1.029	–1.313	1.502	–0.107	1.015	–1.375	1.005	0.318
Italy	1.036	–0.353	1.688	0.195	1.009	–0.165	1.063	0.201
Japan	1.003	0.499	**0.790**	**2.773**	1.004	–0.980	**0.785**	**3.833**
Korea	**0.998**	1.220	1.200	0.165	**0.989**	1.460	**0.996**	0.042
Netherlands	1.022	–0.358	1.416	**3.684**	1.016	–0.283	1.289	**4.366**
Norway	1.027	–0.555	1.004	0.471	1.018	–0.858	**0.822**	1.165
Spain	1.036	–0.389	1.889	0.111	1.007	–1.393	1.150	0.189
Sweden	1.024	0.019	**0.874**	1.110	1.011	–0.377	**0.677**	**2.143**
Switzerland	1.003	0.633	**0.619**	**2.002**	**0.996**	0.926	**0.604**	**1.589**
Average	1.016		1.118		1.006		0.889	

Notes: See notes to table 6.5a.

Theil *U*-statistic is greater than 1, indicating that the RMPSE for the model was greater than for the random walk. We can understand this, taking the particular example of the PPP model for Sweden: in this case, there is a very high positive correlation between the 16-quarter change in the exchange rate and the 16-quarter-ahead forecast from the structural model (around 0.67). The structural forecast, however, overshoots by a long margin—that is, the forecast points in the right direction (tends to be positive when subsequent 16-quarter change is positive, negative when the 16-quarter change is negative), but is much too large in absolute value. Because the forecasted change is too large, the model has a large MSPE and the *U*-statistic is greater than 1. But because it is highly correlated with the actual change in the exchange rate, there is strong evidence against the random walk.

Table 6.7 summarizes the results presented in detail in tables 6.5 and 6. 6. Overall, our conclusion is that the models estimated using panel

Table 6.7
Summary of Forecasting Results

	Panel				Single equation			
	vs. random walk		vs. random walk w/drift		vs. random walk		vs. random walk w/drift	
	u(1)	u(16)	u(1)	u(16)	u(1)	u(16)	u(1)	u(16)
Monetary								
No. of countries in which u() < 1	6	11	13	16	3	1	10	4
Mean value of u()	1.013	0.945	1.004	0.765	1.011	1.460	1.001	1.182
PPP								
No. of countries in which u() < 1	7	14	15	17	2	4	11	8
Mean value of u()	1.004	0.874	0.995	0.701	1.010	1.328	1.000	1.073
Taylor rule								
No. of countries in which u() < 1	2	5	14	16	2	6	2	12
Mean value of u()	1.008	1.142	0.999	0.917	1.016	1.118	1.006	0.889

Notes: This table summarizes results presented for each of the 18 currencies in tables 6.5a–6.5c (in columns headed "Panel) and 6.6a–6.6c (in columns headed "Single equation"). See notes to table 6.5a for details. The "u(1)" columns present results for 1-quarter-ahead forecasts. The "u(16)" columns present results for 16-quarter-ahead forecasts. When u(1) < 1 or u(16) < 1, the panel or single-equation forecast had lower root mean squared prediction error than did the random walk or random walk with drift model.

techniques and producing forecasts at long horizons have some power to beat the random walk. The results are not uniformly strong across the models, in keeping with the EW05 theorem, but it does appear that this approach increases the forecasting power of the models.

6 Conclusions

For many years, the standard for evaluating exchange rate models has been out-of-sample fit. In particular, exchange rate models have been deemed successful or unsuccessful based on their ability to produce better forecasts than the random walk model.

However, many of the models actually imply that the exchange rate should nearly follow a random walk. We should not expect the models to have much power to forecast changes in exchange rates. This might be disappointing news for forecasters working on Wall Street, but may be good news for open-economy macroeconomists. The usual finding-that the models cannot beat a random walk, especially at short horizons—does not mean that open-economy models have been using the wrong exchange rate model, or models that have been refuted by the data.

In fact, and in contrast to many previous studies, we have found evidence that the monetary models do help to forecast changes in exchange rates. That is, with panel techniques, our findings are that the model generally produces better forecast than the random walk—lower mean-squared prediction errors. We find these results encouraging, because they confirm the positive findings of previous studies that have used monetary exchange rate models estimated by panel techniques to forecast exchange rates. But we do not intend to rely on these positive results as our sole piece of evidence in favor of the models. As we have explained, out-of-sample forecasting power relative to the random walk is an unreliable gauge for measuring the models.

At the same time, failure to predict changes in the exchange rate is not a badge of honor for a model. We must develop other means for evaluating and comparing models. This task is made especially difficult because of the fact that plausible models capture the fact that exchange rates are asset prices and are therefore primarily driven by expectations. Because we do not have a direct measure of the market's expectation, we cannot test directly the models.

The array of evidence presented in this paper—including out-of-sample forecasting comparisons—lends weight to monetary models of

exchange rates. That is, to a large extent, the evidence is consistent with the view that nominal exchange rate movements are monetary phenomena. It is, however, especially important to pay attention to how monetary policy is formulated, because expectations of future monetary conditions play an important role in determining current exchange rates.

We do not intend to claim that the evidence in this paper is conclusive in favor of macroeconomic models of the exchange rate. There may well be room for models with private information (such as Bacchetta and van Wincoop [2006]), or models based on the microstructure of foreign exchange markets (e.g., Evans and Lyons [2003]) to improve our understanding of currencies. Moreover, our empirical verification of these models is by no means the final word. We hope, instead, that this paper sparks a renewed interest in the empirical examination and comparison of exchange rate models, but using tests that can fairly support or reject one model in favor of another.

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Endnotes

1. It is important to recognize that the interest semielasticity of money demand depends on the units in which interest rates are expressed. For example, a value of λ of 0.025 estimated with quarterly data when interest rates are annualized and expressed in percent terms must be multiplied by 400 to get the relevant estimate when interest rates are in the same units as the change in the log of the exchange rate.

2. Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Great Britain, Greece, Italy, Japan, Korea, the Netherlands, Norway, Spain, Sweden, and Switzerland.

3. We note that there are missing values for Japan and Norway for October 2001 and for Norway and Switzerland for April 1997 to October 1998.

4. The source for consumer price data used to construct the real exchange rates is the same as in the rest of the paper (IFS, line 64). We use the annual average exchange rate (IFS, line rf) for the nominal exchange rate.

5. See also Mark and Sul (2002) for a related discussion of the asymptotic power of long-horizon regression tests in error-correction models.

6. The results are not reported but are available upon request.

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Comment

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Engel, Mark, and West (EMW) have produced an extremely clear exposition of the recent literature on empirical exchange rate modeling, which has been enjoying a minor renaissance, not least due to the contributions of these three authors. EMW argue that monetary models of exchange rates are finally gaining empirical traction, and that the dismal results found in the literature during the first three decades of the modern (post-1973) float are gradually giving way to more positive ones. Whereas EMW clearly intend to persuade the reader that empirical exchange rate models are back, they deserve credit, too, for emphasizing the still very tentative nature of the new results. These include, not least, the potential sensitivity to sample period that has so plagued earlier attempts to provide robust evidence in favor of structural exchange rate models. There are some important ideas and impressive advances here, but evaluating success is a matter of perspective. Is the glass 10 percent full or 90 percent empty?[1]

1 Noncontroversial Successes of Empirical Exchange Rate Models at Very Long and Very Short Forecast Horizons

Standard textbook empirical exchange rate models have two long-standing areas of success to their credit, about which (for the moment) there is relatively little controversy. EMW present new results to emphasize and strengthen these points. A small number of studies have found that canonical monetary models appear to be able to robustly explain some fraction of long-run movements in the nominal exchange rate. Mark (1995) was the first to formally demonstrate this traction, although Meese and Rogoff (1983b) also found that, at forecasting horizons over two years, the models appeared to outperform the random

walk model over a wide range of parameter grids. Especially in the case of monetary models, it seems plausible to suppose that there must be a link between their ability to forecast long-term nominal exchange rates and the PPP literature, which finds that shocks to the *real* exchange rate tend to damp out only very slowly, with a half-life of two to four years (e.g., Rogoff 1996).[2]

Perhaps the most concrete success for the models comes at very high frequencies. There is now substantial evidence that monetary models, possibly incorporating Taylor rule reaction functions, capture the *qualitative* reaction of exchange rates to central bank interest rate announcements. Thus, for example, when news arrives that raises market expectations about the future path of short-term interest rates (such as a central bank interest rate announcement), a country's currency tends to appreciate. However, both theoretically (e.g., according to plausible parametrizations of either the Dornbusch [1976] model or a modern new open-economy macroeconomics model) or empirically (e.g., Reifschneider, Tetlow, and Williams 1999), interest rate shocks seem to be drowned out by other factors at intermediate time intervals, say from one month to one year.

2 The Big Challenge: One Month to One Year Horizons

The real debate now is over how well structural models can perform at intermediate forecast horizons of one month to one year, arguably the most relevant for policy purposes. Up until now, the main finding in the literature is that it is very difficult to find empirical models that work on a *consistent and robust* basis, even against the apparently weak benchmark of the random model, as Meese and Rogoff (1983a, b) first demonstrated. EMW do not claim to decisively overturn this result, but they do give a wide variety of evidence aimed at raising one's confidence in the relevance of monetary models. It should be noted that the Meese-Rogoff result applies to major floating currencies (e.g., the yen, pound and euro against the dollar) and not necessarily to emerging-market currencies where interest and inflation differentials are often much larger. Also, even among developed economies, Chen and Rogoff (2003) show that there is an important exception for what they term "commodity currencies," such as the Australian, New Zealand, and Canadian dollars. These are countries where commodities form a significant component of the country's overall export basket, and where the world price of the coun-

try's major commodity exports appears to be highly correlated with exchange rate changes, even in out-of-sample fit exercises.

3 A Brief Review of the Meese-Rogoff Result

Given that a core focus of EMW's paper is to show that exchange rate models outperform the naïve random walk model, it is perhaps useful to review Meese and Rogoff (1983a, b) where this benchmark was first established. Meese and Rogoff's paper was published only a decade after the start of the modern floating exchange period, in 1973. (We need not concern ourselves here with how the earlier Bretton Woods system evolved, as post-World War II capital controls were gradually ended across Europe and Japan.) Early research by Dornbusch, Hooper, Morton, Frenkel, Bilson, Frankel and others suggested that simple structural monetary models could do a pretty good job explaining exchange rate movement under floating, at least on an in-sample basis. The typical model included variables such as long- and short-term interest rates, outputs, money supplies, inflation rates, and trade balances (and perhaps other real shocks). For simplicity, the models were typically specified as the difference between home and foreign variables, although most research, including Meese and Rogoff's, did not find this constraint to be a major issue. The canonical empirical equation in the literature could typically be written in semi-reduced form as:

$$s = a_0 + a_1(m - m^*) - a_2(y - y^*) + a_3(i - i^*)$$
$$+ a_4(\pi^e - \pi^{e*}) + a_5 \sum(TB - TB^*) + u$$

where "*" denotes the foreign country, s, m, and y are the (log) exchange rate, money supply, and output, i is the short-term nominal interest rate, π is the expected inflation rate, TB is the trade balance (or sometimes cumulated trade balance), and u is a (possibly serially correlated) error term. The main point of departure in Meese and Rogoff was an effort to test the models out of sample, on the grounds that in-sample tests may obscure problems such as parameter instability and model misspecification.

Meese and Rogoff proceeded by using several alternative estimation methodologies, including OLS rolling regressions, independent variable (IV) rolling regressions, and imposing a grid of parameter restrictions

(the latter approach is developed comprehensively in Meese and Rogoff [1983b]). In each case, the model parameters were estimated using data available up to the initial forecast period. The forecasts themselves, however, were generated using actual realized values of the explanatory variable. Remarkably, the models failed to outperform the random walk model, despite the advantage of being supplied with ex post information. Meese and Rogoff (1983b) found this result to be true across a wide grid of imposed parameters as long as the model predictions were constrained to conform with theoretically imposed signs. Meese and Rogoff considered both mean square forecast error as well as mean absolute forecast error (a nuance forgotten in much of the subsequent literature, but an important one, given evidence the distribution of exchange rate changes typically appears to have fat tails). They also considered a number of variations of the models, including, for example, using price levels to substitute out for money supplies in the core specification to immunize the models from instability in money-demand equations.

Importantly, Meese and Rogoff also experimented with allowing for lagged adjustment and, notably, serial correlation in the error terms. That is, they allowed (say) the one-year-ahead exchange rate forecast to reflect not only actual realized ex post values of the explanatory variables (the "fundamentals" in EMW's terminology), but also projected changes in u, the serially correlated error. In addition to structural models, Meese and Rogoff also looked at a wide variety of time series models, ranging from vector autoregressions to a random walk model with drift. Incredibly, no model seemed to outperform the random walk model across a wide variety of robustness checks. Meese and Rogoff did not actually prove that the random walk model was a *better* forecaster of exchange rates, only that it was not worse; the root mean square errors (RMSEs) for the random walk model were almost never larger across the many experiments. Meese and Rogoff also found that the random walk model outperformed the forward rate (a version of what later became known as "the forward premium" puzzle), but this result is not central to the main point, which is the poor performance of structural models out of sample. Meese and Rogoff's main positive result, as already noted, is that the models appeared to perform better than the random walk at very long time horizons (over two years).

Although these extremely negative results had the salient effect of bringing discipline to the exchange rate literature, it has certainly come as a surprise that they remain so robust some 25 years later. For one thing, Meese and Rogoff had less than ten years of data to work with.

Also, during the time period they examined, financial markets were still evolving. Japan's money markets, for example, were still highly regulated. Even the United States still had interest rate restrictions on consumer checking and savings accounts during most of the time period. At the same time, monetary policy was extremely volatile, making it difficult for markets to anchor expectations in the transition to floating exchange rates. Given these limitations, the presumption was that the models would begin to do better as financial markets developed and as more data accumulated.

Yet, even as more data have become available, the basic empirical result has continued to hold, and is noted in many studies, including those cited by EMW and Rossi—for example, the Frankel and Rose (1995) chapter for the *Handbook of International Economics.* Even the post-1984 "Great Moderation," which has also led to slightly less volatile exchange rates, does not appear to have made them all that much easier to explain.

4 Are Out-of-Sample Tests Too Harsh?

In EMW's argument, out-of-sample fit (where models continue to perform poorly) is at best too harsh a criterion, and possibly inappropriate. They give two arguments, one of which seems easily finessed. The other has been dealt with in the literature but is nevertheless potentially significant. The weaker of EMW's arguments is their claim that there is a certain arbitrariness to out-of-sample fit because root mean forecast error can be reduced by (what they claim) is an innocent transformation of the model. In particular, suppose the exchange rate s is governed by

$$s = x + u$$

where x is the fundamental variable and u is the error term. EMW point out that this simple model can just as well be written as, say,

$$s = \frac{2}{3}s + \frac{1}{3}(x + u).$$

Then if one substitutes in realized values for the explanatory variables (now s and x, instead of just x) the standard error of the resulting equation is only one third as large as in the original equation (since two-thirds of the exchange rate is simply being explained by itself). If I am interpreting this correctly, it would strike me that this "flaw" in using out-of-sample fit as a criterion for exchange rate models would be better described as a "potential pitfall," which the vast majority of studies manage to avoid.

To be fair, EMW give a more subtle version of the same problem in which, after substituting out for the short-term interest rate differential, the forward exchange rate appears on the right-hand side of the estimating equation, instead of the spot exchange rate. Since, empirically, forward and spot rates move almost one for one in the data (as interest rate differentials are small compared to exchange rate volatility), this amounts to virtually the same sleight of hand. Again, it would seem that this problem would be easily avoided, although perhaps some future application will appear where the pitfall is less obvious.

Although they give it less emphasis, EMW's second critique of out-of-sample fit, originally due to Rossi (2005), is more substantive, although it certainly does not really imply dismissing out-of-sample tests. Rossi notes that if the error term u in the exchange rate forecasting equation is serially correlated, then the models may perform poorly when their forecasts fail to take this into account, even when ex post realized values of the explanatory variables are used to construct forecasts. As EMW note, Meese and Rogoff (1983a, b) do in fact report experimenting extensively with incorporating serial correlation into their forecasts. Although they report that these experiments did not overturn their results, one must bear in mind that Meese and Rogoff had only ten years worth of data. Presumably, with a long enough data set, allowing for serial correlation would make a bigger, and potentially decisive, difference.

Ultimately, there is no question that the aim is to have models that can forecast out of sample (as opposed to simply fit out of sample), as well as models that are useful for policy analysis. As Rossi notes in her comments here, out-of-sample tests (of fit and forecasting) will likely remain a fundamental reality check on the success of empirical exchange rate models for long into the future.

5 Do Rationales for the Near-Random Walk Behavior of Exchange Rates also Explain Why a Model May Fit Poorly Out of Sample?

The distinction between out-of-sample fit and out-of-sample forecasting is also relevant in interpreting EMW's very clever observation, based on Engel and West (2005), which is that the exchange rates may exhibit near-random walk behavior even if the underlying fundamentals exhibit substantial serial correlation. The key relationship is their equation (8), where the exchange rate depends on present and future expected values of the fundamentals. Engel and West show the discounting rela-

tionship between the exchange rate and the future fundamentals, implying that the exchange rate may exhibit surprisingly-near random walk behavior, even when the fundamentals damp out at a moderate rate over time. Thus, they argue, one should not be surprised that the exchange rate approximates a random walk, and this should not be interpreted as any kind of evidence against the monetary models. As EMW generously note, their equation (8) is very closely related to an isomorphic equation in Obstfeld and Rogoff (2003), reproduced below:

$$s_t = \frac{\bar{i}\varepsilon}{1+\bar{i}\varepsilon}\sum_{j=t}^{\infty}\left(\frac{1}{1+\bar{i}\varepsilon}\right)^{j-t} E_t\left(m_j - m_j^* + \frac{\rho_j - \rho_j^*}{\bar{i}\varepsilon}\right).$$

Obstfeld and Rogoff (whose derivation is based on a micro-based New Open Economy Macroeconomics Model) argue that for plausible values of the steady-state interest rate i and the interest semielasticity of money demand ε, the discount factor is near 1, and thus relatively small shocks to the risk premium ρ can potentially have very large and long-lasting effects on the exchange rate. (In their model, the risk premium is derived endogenously.) Since risk premia shocks are likely a major source of exchange rate fluctuations, this is potentially an important explanation of exchange rate volatility. Engel and West's explicit calibrations take this point much further and make it more convincing.

The fact that the exchange rate can follow near random walk behavior when governed by mean-reverting monetary fundamentals just might explain why the latter seem of so little use in predicting exchange rates, at least over the relatively short samples international economists typically have to work with. It does not, however, explain why the models fit so badly out of sample using actual realized values of the explanatory variables. Take equation (9) from EMW,

$$(1-b)(s_t - \mathbf{a}'\mathbf{x}_t) = b\rho_t + b(i_t - i_t^*),$$

where (again) ρ is the unobserved risk premium and $\mathbf{a}'\mathbf{x}_t$ are the observed monetary fundamentals. This equation is an example of the type simulated by Meese and Rogoff, discussed above. In this formulation, which is standard in the out-of-sample testing literature, there is no forward-looking sum because expectations are already embodied in the interest differential $i - i^*$. If the model fits poorly out of sample, it must be attributable to parameter instability, misspecification, or some other problem, not to discounting.

Rossi discusses another very intriguing and clever result from Engel and West (2005) that is also applied here. The authors show that even if

fundamentals are of only marginal use in predicting the exchange rate (because of near random walk behavior), it is possible that the exchange rate might be useful in predicting the fundamentals, since the exchange rate embodies expectations about their future evolution. This is a tremendously constructive result that promises to open up a range of new applications. If the relationship is strong, however, it is not clear why it does not show up more strongly in tests of out-of-sample fit. Certainly, if one looks at the one known major exception to the Meese-Rogoff result (the commodity currencies), there is a strong link between commodity currency exchange rates and world commodity prices.

6 Panel Tests and Advances in Out-of-Sample Forecasting Tests

Last but not least, EMW use panel estimation techniques to test the model's out-of-sample forecasting ability. As I have already stated, the authors' results are broadly consistent with the previous literature; there appears to be forecasting power at very long horizons but not at short horizons. It is worth noting that the Meese-Rogoff literature suggests that *in practice* the simple random walk model is a much tougher alternative to beat than the random walk with drift. Most likely the same parameter instability that plagues the structural models also plagues the random walk with drift model. The final tables of the present paper confirm this well-known empirical result; the models perform much less impressively when compared against the simple random walk model than when compared against the random walk with drift.[3] It is also worth noting that the table reveals some differences across tests for whether the monetary models improve on the random walk model; in some cases, Theil's U-statistic is greater than 1 and yet the Clark-West test is significant. The latter, of course, is a state-of-the-art test based on the premise standard comparisons are biased in favor of the random walk model (since there are no parameters to be estimated) and contains an adjustment for that bias that favors the models with fundamentals. One wonders whether there may be cases where the adjustment (which is asymptotically correct) may prove to be too large, although further applications and study should resolve this lingering doubt over time.

7 Concluding Remarks

This is a very nice paper that synthesizes the successes of a valuable research program. But despite important methodological improvements,

and a number of clever new ideas, exchange rates remain a very tough nut to crack, even after the Great Moderation in macroeconomic variables. Right now, things still look pretty good if we can call the glass 10 percent full.

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Endnotes

1. The continuing difficulties in understanding exchange rates have important implications not only for policymakers but also for economic theorists. For example, there has been a large recent literature on explaining speculative attacks in fixed exchange rates. Although there are a couple notable exceptions, most of this literature assumes a clear and stable connection between fundamentals in the post-attack floating exchange rate. But a great body of evidence weighs against this assumption. As Flood and Garber (1984) first illustrated (see also Obstfeld and Rogoff 1996), if one cannot tie down the post-attack floating rate, one cannot tie down the timing of the attack either. Difficulties in explaining floating rates may indeed be the root of instability, even under fixed rates.

2. EMW endorse the Benigno (2004) explanation of the purchasing power parity puzzle, where very high correlation in the error term on the Taylor rule, interacting with fixed Calvo contracts, can lead to very long half-lives for deviations from PPP. This explanation, too, is very clever, although one wonders whether the Calvo price-setting parameters should be thought of as fixed regardless of the Taylor rule, especially when serial correlation in the error term is producing such long-lived real exchange rate effects.

3. The first paper to implement the panel approach, Mark and Sul (2001), used the random walk with drift model on logical a priori grounds; the newer results here suggest that this approach may overstate the performance of the structural models.

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Comment

Barbara Rossi

1 Reviewing Engel and West's (2005) Main Result

Let's focus on a simple monetary model of exchange rates (Obstfeld and Rogoff [1996]):

$$E_t(s_{t+1} - s_t) = -\frac{1}{\eta}(f_t - s_t). \tag{1}$$

where f_t is the fundamental and s_t is the nominal exchange rate (for simplicity, we let the fundamental be a scalar). Let $b \equiv [\eta/(1 + \eta)]$ denote the discount factor; the no-bubble solution is:

$$s_t = (1 - b)\sum_{s=0}^{\infty} b^s E_t f_{t+s} \equiv F_t.$$

The literally has tested these models by using either in-sample (e.g., OLS/GMM estimation of [1]), our out-of-sample methods. While the in-sample evidence tends to find significant fit, out-of-sample analyses usually find that the model performs worse than a random walk: $E_t(s_{t+1} - s_t) = 0$. The latter influential result was first discovered by Meese and Rogoff (1983a, 1988) and has become known as the "Meese-Rogoff puzzle."

In their important paper, Engel and West (hereafter EW) (2005) claim that it is not surprising that the random walk performs better, as this is exactly what the model predicts when $b = 1$. To understand this claim, consider the following data-generating process:

$$\begin{cases} \Delta f_{t+1} = \rho \Delta f_t + \varepsilon^f_{t+1} \\ s_{t+1} = -(1/\eta) f_t + (1 - 1/\eta) s_t + \varepsilon^e_{t+1} \end{cases},$$

$$\text{where} \begin{pmatrix} \varepsilon^f_{t+1} \\ \varepsilon^e_{t+1} \end{pmatrix} = \begin{pmatrix} 1 & \sigma_{fe} \\ 0 & 1 \end{pmatrix} \begin{pmatrix} \eta^f_{t+1} \\ \eta^e_{t+1} \end{pmatrix} \tag{3}$$

$\varepsilon^e_{t+1} = -E_t(s_{t+1} - s_t)$ from (1), $F_{t+1} \equiv (1-b)\Sigma_{s=0}^{\infty} b^s E_t f_{t+s}$ and $cov(\varepsilon^f_{t+1}, \varepsilon^e_{t+1}) = \sigma_{fe}$. EW (2005) assume $\sigma_{fe} = 0$. Also, let $\Phi(L) \equiv (1-\rho L)$ be the autoregressive polynomial for the fundamentals, with L being the lag operator, so that (3) is equivalent to $\Phi(L)\Delta f_{t+1} = \varepsilon^f_{t+1}$. By standard manipulations, if $|b\rho| < 1$: $\Delta F_t - E_{t-1}\Delta F_t = \Sigma_{j=0}^{\infty} b^j (E_t \Delta f_{t+j} - E_{t-1}\Delta f_{t+j}) = (1-b\rho)^{-1}\varepsilon^f_t$ and $E_{t-1}\Delta F_t = (1-b)E_{t-1}\Sigma_{j=0}^{\infty} b^j \Delta f_{t+j} = (1-b)\rho(1-b\rho)^{-1}\Delta f_t$. Thus, when the model is true ($s_{t+1} = F_{t+1}$, i.e. $\Delta s_{t+1} = \Delta F_{t+1}$), $\Delta s_t = (1-b\rho)^{-1}\varepsilon^f_t + (1-b)\rho(1-b\rho)^{-1}\Delta f_t$, so that:

$$\Delta s_{t+1} = \Phi(b)^{-1}\, \varepsilon^f_{t+1} \text{ for } b \approx 1 \tag{4}$$

where $\Phi(b) \equiv (1-\rho b)$. Equation (4) implies EW's (2005) important result that, even though exchange rates are determined in equilibrium as the net present value of fundamentals, nevertheless exchange rate changes are unpredictable given information at time t, when the discount factor b is sufficiently high. Note that the proposition says that exchange rates are unpredictable given information at time t, not that they are random walks completely unrelated to the fundamentals (in fact, σ_{fe} could be different from zero).

2 Empirical Evidence in Favor of EW's (2005) Conjecture

To substantiate their claims, EW (2004, 2005) report empirical evidence on variance bounds and Granger-causality tests. In what follows, we replicate their empirical findings by using the same database as in EW (2004). The database consists of quarterly data from 1974:1 to 2003:1 for bilateral exchange rates of six major G7 countries relative to the U.S. dollar.

2.1 *Implications for Variance Ratios*

How much of the volatility of exchange rates can the model explain? EW (2004) note that (4) implies that the variance of Δs_t explained by the net present value of the fundamentals is $var(\Delta F_t) = var\,[\Phi(b)^{-1}\varepsilon^f_t]$. To estimate the latter, they estimate an autoregression for the fundamentals: $\Phi(L)\Delta f_t = \alpha + \varepsilon^f_t$, where $\Phi(L) = 1 - \Sigma_{j=1}^{p}\Phi_j L^j$, using $p = 4$. The estimates of $\Phi_1, \ldots, \Phi_4$ and of the time series of the fitted residuals ε^f_t together with

Table 6C2.1
Estimated Variance Ratios (i.e., Fraction of the Variance of Exchange Rates Explained by Fundamentals)

Δf:	Canada	France	Italy	Germany	Japan	U.K.
$\Delta(m-m^*)-\Delta(y-y^*)$	**1.212**	**0.352**	**0.373**	**0.345**	**0.446**	**0.644**
$\Delta(p-p^*)$	0.211	0.186	0.134	**0.422**	0.090	0.298
$\Delta(p-p^*)-\Delta(i-i^*)$	0.198	0.145	0.100	**0.343**	0.025	0.283
$\Delta(m-m^*)$	**1.447**	**0.423**	**0.703**	0.173	**0.388**	**0.686**
$\Delta(i-i^*)$	0.005	0.003	0.001	0.002	0.001	0.002
$\Delta(y-y^*)$	0.097	0.040	0.144	0.035	0.069	0.024

Note: Bold indicates fractions above 30%.

a calibrated value of b allow them to estimate the contribution of the variability of the net present value of fundamentals to the total variability of exchange rates: *Variance ratio* $\equiv var[\Phi(b)^{-1}\varepsilon_t^f]/var(\Delta s_t)$. Table 6C2.1 reports variance ratio estimates by calibrating $b = 0.99$ and using the same fundamentals as in EW (2004): money/income differentials $[\Delta(m-m^*)-\Delta(y-y^*)]$, price differentials $[\Delta(p-p^*)]$, and price/interest rate differentials $[\Delta(p-p^*)-\Delta(i-i^*)]$. We also experimented with other fundamentals considered in EW (2005), including money differentials $[\Delta(m-m^*)]$, interest rate differentials $[\Delta(i-i^*)]$, and income differentials $[\Delta(y-y^*)]$. Table 6C2.1 replicates EW's (2004) finding that when $b \approx 1$ the variability of fundamentals explain about 40 percent of the variability of exchange rates. Indeed, the fraction of variability of exchange rates that the procedure attributes to fundamentals is high (about or above 40 percent) for most fundamentals, with the exception of interest rate and income differentials.[1]

- *EW (2004, p. 124) therefore conclude that: "models in which the exchange rate is . . . a discounted sum of . . . future fundamentals can account for a sizeable fraction of the variance of . . . exchange rates when the discount factor is large. . . ."*

2.2 *Implications for Granger Causality Tests*

Equation (2) implies that exchange rates must Granger-cause fundamentals. Table 6C2.2 replicates EW's (2005) finding that exchange rates Granger-cause a variety of fundamentals (panel A) and that the reverse is not true (panel B). The table reports p-values of the Granger-causality tests—numbers below 0.05 imply evidence in favor of Granger causality[2].

Table 6C2.2
Bivariate Granger Causality Tests, Different Measures of Δf_t

Δf:	Canada	France	Italy	Germany	Japan	U.K.
A. p-values of $H_0: \beta_0 = \beta_1 = 0$ *in* $\Delta f_{t+1} = \beta_0 + \beta_1 \Delta s_t + \gamma \Delta f_t$						
$\Delta(m - m^*)$	0.10	**0.01**	**0**	0.63	0.31	**0.03**
$\Delta(p - p^*)$	0.82	0.67	**0**	**0**	**0**	**0**
$\Delta(i - i^*)$	0.21	0.15	0.42	0.27	**0.01**	**0.01**
$\Delta(m - m^*) - \Delta(y - y^*)$	0.22	**0.01**	0.40	**0**	0.31	**0.01**
$\Delta(y - y^*)$	0.31	**0.10**	0.28	0.15	1	0.12
$\Delta(p - p^*) - \Delta(i - i^*)$	0.88	**0**	**0**	**0**	**0**	**0**
B. p-values of $H_0: \beta_0 = \beta_1 = 0$ *in* $\Delta s_{t+1} = \beta_0 + \beta_1 \Delta f_t + \gamma \Delta s_t$						
$\Delta(m - m^*)$	0.16	0.21	0.31	0.69	0.38	0.69
$\Delta(p - p^*)$	0.23	0.32	0.33	0.67	0.51	0.36
$\Delta(i - i^*)$	0.25	0.91	0.25	1	0.75	0.71
$\Delta(m - m^*) - \Delta(y - y^*)$	0.21	0.17	0.12	0.78	0.51	0.50
$\Delta(y - y^*)$	0.11	0.72	0.31	0.85	0.31	0.55
$\Delta(p - p^*) - \Delta(i - i^*)$	0.32	0.46	0.37	0.35	0.77	0.48

Note: Bold denotes p-value lower than 5%.

- *EW (2005) therefore conclude that, even though there is no evidence that fundamentals Granger-cause future exchange rates, there is substantial empirical evidence that exchange rates Granger-cause fundamentals, in accordance with the net present value models of nominal exchange rates.*

2.3 *Implications for Out-of-Sample Forecasts*

In order to provide further evidence that fundamentals help predict exchange rates, Engel, Mark, and West report forecasts based on long-horizon panel regressions. They argue that fundamentals and exchange rates are cointegrated if one uses panel unit root tests. They also show that, if one exploits the panel dimension, fundamentals are able to provide better forecasts of future exchange rates than a simple random walk model, especially at long horizons.

3 Robustness Checks

It is important to realize that EW's (2004, 2005) and EMW's results rely on a variety of assumptions:

1. An autoregression with 4 lags captures well the behavior of fundamentals. But how well do 4 lags of the fundamentals forecast future values of fundamentals?

2. The variance ratios are precisely estimated. But how much uncertainty is there around the point estimates of the variance ratios?

3. The variance ratios are informative about the contribution of fundamentals in explaining the variability of exchange rates. But how informative are variance ratios?

4. The Granger-causality tests are performed over the full sample (1973–2003), therefore assuming stability in both the exchange rates and the fundamentals in the last three decades. However, that period witnessed a variety of structural changes in the G7 countries (e.g., changes in monetary policy in the United States [Clarida, Galí, and Gertler 2000], the introduction of a common currency in Europe). Are the Granger-causality results robust to instabilities?

5. The long-horizon regressions are valid only if the nominal exchange rates and the fundamentals are cointegrated. Is the empirical evidence in favor of cointegration robust?

6. The long-horizon panel regression forecasts are evaluated by using tests of equal out-of-sample forecasting ability. Are the results of such tests robust to the lack of cointegration?

In what follows, we check the robustness of the results to each of these assumptions.

3.1 Lag Length and Fit of the Autoregression

A detailed analysis of the AR(4) model in EW (2004) shows that only a few of the estimates are statistically different from zero (a Bayesian Information Criterion [BIC] would pick one lag for all countries as the optimal lag length). We therefore checked whether results in table 6C2.1 were robust to setting the order of the autoregression to 1. Table 6C2.3 reports the results. Overall, the estimates of the variance ratios drop. For example, when the money/income ratio is used as the fundamental, it drops from 40 percent to between 20 and 30 percent.

• *We therefore conclude that, although the actual fraction of variance of exchange rates explained by fundamentals is sensitive to the underlying dynamics of the fundamentals, the results are very robust, as this fraction is still sizeable.*

Table 6C2.3
Variance Ratios: Robustness to the Order of the Autoregression

Δf:	Canada	France	Italy	Germany	Japan	U.K.
$\Delta(m-m^*)-\Delta(y-y^*)$	**0.865**	0.181	0.190	0.216	**0.300**	**0.327**
$\Delta(p-p^*)$	0.158	0.050	0.041	0.145	0.022	0.094
$\Delta(p-p^*)-\Delta(i-i^*)$	0.153	0.031	0.038	0.073	0.012	0.097
$\Delta(m-m^*)$	**0.899**	0.195	0.246	0.126	0.291	**0.375**
$\Delta(i-i^*)$	0.006	0.008	0.002	0.009	0.003	0.006
$\Delta(y-y^*)$	0.097	0.031	0.073	0.037	0.043	0.035

Note: Bold indicates fractions above 30%.

3.2 *How Much Uncertainty Is There around EW's (2004) 40 Percent Point Estimate?*

To evaluate the estimation uncertainty around the variance ratio statistic, we calculate confidence intervals. Since the Monte Carlo empirical distribution of the estimated variance ratio statistic is highly asymmetric, the normal approximation and the use of the mean as a point estimate may not be appropriate. We thus use Kilian's (1998) bias-adjusted bootstrap, setting $b = 0.99$ and the autoregressive lag length equal to 4. Table 6C2.4 reports median unbiased estimates of the variance ratio as well as 90 percent confidence intervals in parentheses. Note that, due to the asymmetry, the point estimates in Table 6.8 are a poor approximation of the median: median estimates are significantly higher. Note also that the confidence intervals are really wide.

- *We therefore conclude that EW's (2004) point estimates actually underestimate the true variance ratios, therefore providing additional evidence in favor of their conjecture. However, there seems to be substantial uncertainty around such estimates, which generally can be anywhere between 10 percent to above 100 percent.*

3.3 *Do the Dynamics in the Net Present Value of Fundamentals Explain Exchange Rates?*

How informative are variance ratios? How should we interpret the result that fundamentals explain about 40 percent of exchange rate variability? We generate independent white noise fundamentals with variance equal to the estimated variance of the fundamentals. The estimated

Table 6C2.4
90 Percent Confidence Intervals for the Variance Ratios, Killian's (1998) Bootstrap

Canada	France	Italy	Germany	Japan	U.K.
		$\Delta(m-m^*)$			
1.6994	**0.6455**	**1.2579**	0.2207	**0.4865**	**1.0233**
(0.75, 4.10)	(0.16, 3.60)	(0.24, 10.00)	(0.08, 0.65)	(0.17, 1.43)	(0.29, 4.41)
		$\Delta(p-p^*)$			
0.2559	**0.5676**	0.2819	1.1136	0.1799	**0.5182**
(0.09, 077)	(0.05, 11.34)	(0.04, 5.43)	(0.13, 27.62)	(0.03, 2.49)	(0.10, 4.96)
		$\Delta(i-i^*)$			
0.0055	0.0031	0.0017	0.0019	0.0018	0.0025
(.002, .01)	(.001, .005)	(.0008, .003)	(.0008, .004)	(.0008, .003)	(.0014, .004)
		$\Delta(m-m^*)-\Delta(y-y^*)$			
1.3521	**0.5093**	**0.5083**	**0.4585**	**0.5423**	**0.8736**
(0.61, 2.87)	(0.13, 2.25)	(0.15, 1.85)	(0.14, 1.59)	(0.20, 1.48)	(0.30, 3.07)
		$\Delta(y-y^*)$			
0.112	0.0449	0.1814	0.0385	0.0811	0.0262
(0.05, 0.25)	(0.02, 0.10)	(0.05, 0.58)	(0.01, 0.08)	(0.03, 0.20)	(0.01, 0.05)
		$\Delta(p-p^*)-\Delta(i-i^*)$			
0.2392	0.259	0.1546	**0.8603**	0.0341	**0.4381**
(0.09, 0.64)	(0.05, 2.17)	(0.04, 0.81)	(0.10, 26.8)	(0.01, 0.12)	(0.09, 2.88)

Note: bold denotes fractions above 30%.

Table 6C2.5
Variance Ratios with Independent Fundamentals

	Canada	France	Italy	Germany	Japan	U.K.
$p = 1$	0.9540	0.1863	0.2666	0.1006	0.1488	0.0949
$p = 4$	0.7534	0.1916	0.0723	0.500	0.1857	0.0678

variance ratios are reported in table 6C2.5 for a choice of the lag length (p) equal to 1 or 4. The table shows that even completely unrelated fundamentals can explain a sizeable fraction of the variability of exchange rate fluctuations, provided that their variance is similar to that of the net present value of the fundamentals. In other words, variance ratios do not provide much information on the nature of the dynamic relationship between exchange rates and fundamentals—they only provide information on the size of their relative unconditional variances.

Table 6C2.6
Correlations between Δs_t and ΔF_{t+j}

j	–5	–4	–3	–2	–1	0	1	2	3	4	5
Canada	–0.03	0.00	–0.05	0.08	–0.01	0.07	–0.09	–0.16	0.14	–0.19	–0.03
France	0.06	–0.05	0.06	0.14	0.14	0.00	**0.31**	0.01	–0.02	0.16	0.20
Italy	0.06	0.05	–0.11	0.13	0.10	–0.10	0.02	–0.13	0.02	0.03	0.06
Germany	–0.06	0.07	–0.16	0.06	0.06	–0.05	**0.32**	–0.13	0.00	0.10	0.06
Japan	0.23	0.14	0.04	0.00	0.02	0.12	0.13	–0.14	0.02	–0.06	0.02
U.K.	0.05	0.07	–0.07	0.12	0.07	–0.11	–0.07	–0.08	0.01	0.02	–0.02

Note: bold indicates correlations above 20%.

What is, then, the dynamic relationship between fundamentals and exchange rates? Table 6C2.6 shows that there is some correlation but it is sizeable (around 30 percent) only for France and Germany at one lag.

- *We therefore conclude that, although the variance ratios are high, such measures do not capture correlations between exchange rates and fundamentals—and such correlations are small. It is therefore important to complement the analysis with Granger-causality tests, as in EW (2005).*

3.4 *How Robust Are the Granger-Causality Results to Parameter Instabilities?*

Are there parameter instabilities in the data? And how do these affect EW's (2005) results in table 6C2.2? Table 6C2.7 reports p-values of the Andrews (1993) test for parameter instabilities in regressions where exchange rates Granger-cause fundamentals (panel A), and where fundamentals Granger-cause exchange rates (panel B). Values less than 0.05 in the table imply evidence in favor of parameter instabilities at 5 percent significance level. Recall that EW (2005) found evidence that exchange rates Granger-cause fundamentals, but no evidence that fundamentals Granger-cause exchange rates. Table 6C2.7 shows that there is evidence of instabilities exactly in the situations in which EW (2005) find empirical evidence of Granger-causality. In other words, the evidence that exchange rates Granger-cause fundamentals is plagued by instabilities, whereas there is no evidence of parameter instabilities when examining whether fundamentals Granger-cause exchange rates.[3]

In principle, the presence of parameter instabilities could invalidate the empirical evidence in favor of Granger-causality reported in table

Table 6C2.7
Andrews' (1991) QLR Test for Time-Varying Parameters

Δf:	Canada	France	Italy	Germany	Japan	U.K.
A. p-values of stability of (β_{0t}, β_{1t}) *in:* $\Delta f_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \gamma\Delta s_t$						
$\Delta(m - m^*)$	**0**	**0.02**	**0**	**0.02**	**0.09**	**0**
$\Delta(p - p^*)$	0.11	**0**	**0**	**0**	**0**	**0**
$\Delta(i - i^*)$	0.44	**0**	0.12	**0.03**	**0**	**0**
$\Delta(m - m^*) - \Delta(y - y^*)$	**0.05**	0.45	**0**	**0**	**0**	**0**
$\Delta(y - y^*)$	**0**	**0.06**	0.28	**0.06**	**0.05**	1
$\Delta(p - p^*) - \Delta(i - i^*)$	**0.08**	**0**	**0**	**0**	0.21	**0**
B. p-values of stability of (β_{0t}, β_{1t}) *in:* $\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta f_t + \gamma\Delta s_t$						
$\Delta(m - m^*)$	0.43	0.62	0.32	0.66	0.20	1
$\Delta(p - p^*)$	**0.02**	**0**	**0**	0.52	**0**	0.15
$\Delta(i - i^*)$	0.82	0.47	0.56	0.36	**0.09**	0.74
$\Delta(m - m^*) - \Delta(y - y^*)$	0.35	0.54	0.61	0.12	0.58	0.81
$\Delta(y - y^*)$	0.18	0.82	0.51	0.72	0.76	0.16
$\Delta(p - p^*) - \Delta(i - i^*)$	0.20	0.23	0.26	0.73	**0.02**	0.14

Note: bold denotes significant at 10%.

6C2.2, since a maintained hypothesis of Granger-causality tests is that parameters are stable.[4] To assess the robustness of EW's (2005) result—that exchange rates Granger-cause fundamentals, we perform a joint test of parameter stability and Granger-causality by using the test proposed by Rossi (2005a, 2006). Table 6C2.8 reports p-values for the Exp-W* test in Rossi (2005a); p-values less than 0.05 mean that there is evidence that the regressor has some explanatory power, although such a relationship might be unstable over time. The results show that there is ample evidence that exchange rates Granger-cause fundamentals even though their relationship is unstable over time. Interestingly, now we also find (for some fundamentals and some countries) empirical evidence that fundamentals Granger-cause future exchange rates, although, again such causal relationship is unstable over time, and therefore could be very difficult to exploit for forecasting purposes.

• *Therefore, we conclude that EW's (2005) results that exchange rates predict future fundamentals are very robust to the presence of parameter instabilities at unknown times. Interestingly, however, there is some evidence that fundamentals also Granger-cause future exchange rates for some fundamentals and some countries, although such a relationship is unstable over time.*

Table 6C2.8
Optimal Joint Test for Significance and Instability (Rossi [2005])

Δf:	Canada	France	Italy	Germany	Japan	U.K.
A. p-values of $H_0: \beta_t = \beta = 0$ *in* $\Delta f_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \gamma\Delta f_t$						
$\Delta(m - m^*)$	**0.01**	**0**	**0**	**0.03**	0.33	**0.05**
$\Delta(p - p^*)$	0.48	**0**	**0**	**0**	**0**	**0**
$\Delta(i - i^*)$	0.61	**0.02**	0.43	0.17	**0.01**	**0.01**
$\Delta(m - m^*) - \Delta(y - y^*)$	0.14	**0.04**	**0**	**0**	**0.01**	**0.03**
$\Delta(y - y^*)$	**0.07**	**0.02**	**0**	**0.05**	0.44	0.49
$\Delta(p - p^*) - \Delta(i - i^*)$	0.70	**0**	**0**	**0**	**0**	**0**
B. p-values of $H_0: \beta_t = \beta = 0$ *in* $\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta f_t + \gamma\Delta s_t$						
$\Delta(m - m^*)$	0.25	0.38	0.35	1	0.54	1
$\Delta(p - p^*)$	0.18	**0**	**0.04**	0.80	**0.03**	**0**
$\Delta(i - i^*)$	0.77	0.87	0.65	0.84	0.39	1
$\Delta(m - m^*) - \Delta(y - y^*)$	0.20	0.40	0.50	**0**	0.75	1
$\Delta(y - y^*)$	0.24	1	0.62	1	0.21	0.53
$\Delta(p - p^*) - \Delta(i - i^*)$	0.50	0.33	0.27	0.72	**0.06**	**0.03**

Note: bold denotes significant at 10%.

3.5 *How Robust Are Panel Cointegration Results?*

As further evidence in support of the usefulness of economic models of exchange rate determination, EMW report that out-of-sample forecasts of exchange rates at long horizons improve substantially when using economic fundamentals as regressors. They focus on the predictive regression

$$s_{i,t+h} - s_{i,t} = \beta_h z_{i,t} + \varepsilon_{i,t+h}, \tag{5}$$

where $z_{i,t} = x_{i,t} - s_{i,t}$, $x_{i,t}$ is the fundamental, $\varepsilon_{i,t+h} = \alpha_i + \theta_t + u_{i,t}$, i indexes the country, t indexes the time period, and h is the forecast horizon. EMW estimate (5) both with panel regression methods and with univariate VECMs for each country. In both cases, they find substantial improvements in the forecasting ability of models relying on fundamentals at long horizons relative to the random walk. Such improvements are most striking when using PPP fundamentals.

As pointed out by EMW, the predictive regression makes sense when z is a stationary variable (that is when the exchange rate and the fundamentals are cointegrated). This may be problematic in the case of PPP fundamentals, as real exchange rates are very persistent variables, with roots that are statistically indistinguishable from unity (see Rossi 2005c). They report empirical evidence in favor of stationarity based on Sul's

Table 6C2.9
Unit Root Tests and the Robustness of Panel Forecasting Tests

A. Unit root tests on the factors				C. *Panel regressions*		D. *VECM single equation*	
	constant	constant & trend	h	I(1)	I(0)	I(1)	I(0)
F_{1t}	−2.519	−2.615	1	0.054	0.052	0.053	0.044
	B. Pooled unit root tests		2	0.094	0.053	0.066	0.059
	constant	constant & trend	5	0.124	0.062	0.091	0.070
$P^{\tau}_{\hat{e}}$	1.068	1.068	10	0.150	0.079	0.170	0.086
$P^{c}_{\hat{e}}$	2.069*	2.069*	20	0.201	0.113	0.226	0.120

Note: Asterisks denote significance at 5% (5% critical values are: −1.95 in panel A and ±1.64 in panel B). Panels C and D report empirical rejection frequencies of out-of-sample tests of equal forecast accuracy against a random walk. Nominal size is 0.05. Columns labeled I(1) and I(0) denote, respectively, cases where fundamentals and the exchange rates have a unit root and when they are stationary. The forecast horizon is h.

(2006) test. We therefore examine the robustness of EMW's results to other tests for stationarity of z. We consider Bai and Ng's (2004) panel unit root tests for the presence of unit roots in the common factors and in the idiosyncratic components of z_{it}. We focus on the case of PPP fundamentals for which Bai and Ng's (2002) test selects 1 common factor. Unreported results show that the largest roots in the common factors and in the idiosyncratic components are close to unity. Panels A and B in table 6C2.9 show that we cannot reject a unit root in the common factor, nor in the pooled regression tests with a deterministic trend. There seems to be less evidence of nonstationarity in the pooled regressions with a constant. Unreported results for monetary fundamentals show that unit root tests on the idiosyncratic component again do find some nonstationarity, while the pooled tests do not find any nonstationarity.

- *We conclude that there is high persistence in the relationship between exchange rates and PPP fundamentals, and the evidence on cointegration may not be robust to the use of other panel unit root tests, as the results depend on the specification. Results for monetary fundamentals are more robust.*

3.6 *How Robust Are Out-of-Sample Long-Horizon Panel Forecasts?*

The fact that the exchange rate and the fundamentals may not be cointegrated could imply that the long horizon regression results are spurious. However, they might still provide good out-of-sample forecasts, as EMW show. Whether such forecasts are statistically better than the ran-

dom walk's forecast depends on the outcome of tests of forecast comparison. Such tests, however, may have poor properties if the regressors have unit roots.[5] We therefore consider what happens to rejection rates of tests for equal predictive ability in the presence of high persistence.

We consider the following data-generating process:

$$s_{i,t} = \lambda_i' F_t + u_{i,t}, \tag{6}$$

where F_t is a $(r \times 1)$ vector containing the factors $F_{j,t}$, $F_{j,t} = \rho_j F_{j,t-1} + e_{j,t}$, $j = 1, \ldots, r$, and $u_{i,t} = \rho_{i,e} u_{i,t-1} + e_{i,t}$, $i = 1, \ldots N$. We let $\rho_f = 1$ to mimic the realistic situation in which the PPP fundamental has a unit root. We follow EMW and use the Clark and West's (2006) test to evaluate out-of-sample predictive ability.[6] The null hypothesis of the test is that the nominal exchange rates are random walks and the alternative is that the fundamentals help forecasting the nominal exchange rates. To check the properties of the test, we let the nominal exchange rate be uncorrelated with the fundamental ($\lambda_i = 0$) and let the nominal exchange rate be highly persistent ($\rho_{i,e} = 1$). We generated 2,000 Monte Carlo simulations of (6) and forecasted future exchange rates using the same estimation procedures as in EMW, namely panel regressions and single-equation VECM regressions applied to each series in the long-horizon regression model (5). We consider horizons from 1 to 20, which encompass horizons 1 and 16, considered by EMW. Panels C and D in table 6C2.9 report rejection frequencies of nominal 5 percent tests of the null hypothesis that the true model is a random walk. Columns labeled I(1) denote situations in which $\rho_{i,e} = 1$ (similar results hold if the root is not exactly unity but close to unity). As a comparison, we also report rejection frequencies for the stationary case ($\rho_{i,e} = \rho_{i,f} = 0$), labeled I(0). The horizon is denoted by h. It is clear that in the stationary case, rejection frequencies are very close to their nominal value, whereas the tests grossly over-reject the random walk at long horizons in the presence of nonstationarity.

- *We conclude that out-of-sample forecast comparison tests tend to over-reject the null hypothesis of a random walk in the presence of roots close to unity.*

4 How Do We Reconcile These In-Sample Results with Meese-Rogoff's Stylized Fact?

First, is the Meese-Rogoff stylized fact still alive? Table 6C2.10 shows forecast comparison results between models of exchange rate determi-

Table 6C2.10
Tests for Out-of-Sample Forecasting Ability, Different Measures of Δf_t

	Canada	France	Italy	Germany	Japan	U.K.
A. MSE difference between the model: $\Delta f_{t+1} = \beta_{0t} + \beta_{1t}\Delta s_t + \beta_{2t}\Delta f_t$ *and the random walk*						
$\Delta(m - m^*)$	0.75	–1.29***	–0.23***	–0.93***	–1.01***	–1.48***
$\Delta(p - p^*)$	0.97	–0.46***	0.43***	0.28***	–2.16***	0.42***
$\Delta(i - i^*)$	0.98***	2.54	3.47	1.55	–0.13***	2.10
$\Delta(m - m^*) - \Delta(y - y^*)$	0.65	–0.12***	–0.77**	–0.23***	–0.57***	–0.75***
$\Delta(y - y^*)$	1.38	–0.09***	1.92	–0.35***	1.29	–0.27***
$\Delta(p - p^*) - \Delta(i - i^*)$	1.35	2.12	0.70**	–0.10***	–2.65***	0.64***
B. MSE difference between the model: $\Delta s_{t+1} = \beta_{0t} + \beta_{1t}\Delta f_t + \beta_{2t}\Delta s_t$ *and the random walk*						
$\Delta(m - m^*)$	0.60	1.41	1.35	1.45	1.38	1.26
$\Delta(p - p^*)$	1.78	0.68	0.91***	1.02	1.09	1.42*
$\Delta(i - i^*)$	1.03	1.84	0.78**	2.10	0.08	0.87*
$\Delta(m - m^*) - \Delta(y - y^*)$	0.82	1.52	0.67**	1.63	1.93	1.19
$\Delta(y - y^*)$	0.98	2.01	1.16	1.37	1.14***	0.86
$\Delta(p - p^*) - \Delta(i - i^*)$	1.65	1.44	0.97***	1.96	0.49	1.30**

Note: Positive values imply that the model forecasts worse than the random walk. Asterisks denote rejections of the null hypothesis of equal predictive ability at 1% (***), 5% (**), and 10% (*) significance levels.

nation with various fundamentals and the random walk. Parameters are reestimated over time using a rolling forecasting scheme, with a window equal to half of the sample size. The table reports mean square error differences between the model and the random walk, rescaled by a measure of standard deviation.[7] Positive values in the table imply that the random walk forecasts perform better than the model. The table shows that even though the fundamentals may explain a sizeable proportion of the in-sample total variance of exchange rates, they are still not better than a random walk for forecasting exchange rates out of sample. Interestingly, instead, lagged exchange rates not only Granger-cause fundamentals in-sample, but also for some currencies significantly help in forecasting future fundamentals, even out of sample.

Note that the results in table 6C2.10 rely on forecast comparisons of models in which the fundamentals are estimated in first differences. This specification has been chosen here because of consistency with the analysis in tables 6C2.7 and 6C2.8 and EW (2005), although it is different from that used in the original Meese and Rogoff (1983a, 1988) papers. The economic models (in levels) considered by Meese and Rogoff (1983a, 1988) are such that $s_t = \beta f_t + u_t$, where the residuals u_t are typi-

cally serially highly correlated, for example, $u_t = \rho u_{t-1} + \varepsilon_t$, where ε_t is an unforecastable error. Meese and Rogoff (1983a, 1988) estimate the serial correlation to better forecast the exchange rates: $E_t s_{t+1} = \beta E_t f_{t+1} + \rho(s_t - \beta f_t)$. Note that this is somewhat similar to including a lagged cointegrating vector in the model. However, in the data $\rho \approx 1$, and usual estimation procedures typically lead to a downward-biased estimate of such parameter in this case. When comparing a model with a random walk (that imposes $\beta = 0$, $\rho = 1$), the random walk's mean square forecast error could be lower because the bias in the estimate of ρ could more than offset the gain in exploiting the information on the fundamentals if β is not "big." Therefore, the random walk model might even forecast better if the model is the true data-generating process. Rossi (2005b) shows that once parameter estimation error in ρ is taken into account, tests of equal predictive ability do not reject the hypothesis that the model and the random walk are equivalent. If we were sure that $\rho = 1$, then we might impose that in the estimation to improve the forecasting ability of the model. An interesting remark is that Meese and Rogoff (1983b) did allow $\rho = 1$ in their grid search, but this did not improve the model's forecasting ability.[8] Imposing $\rho = 1$ in the estimation would mean estimating the model in first differences. As we have seen throughout this discussion, models estimated in first differences are plagued by parameter instabilities. So that could justify why the model does not forecast better than a random walk, even when the model is estimated in first differences, and is consistent with Meese and Rogoff's (1983b) results.

- *We therefore conclude that Meese and Rogoff's stylized fact is still present in the data, no matter whether the models are estimated in first differences, or with a serial correlation correction, as in Meese and Rogoff's original works.*

Why is there such a big difference between in-sample and out-of-sample performance, then? The previous analysis suggests that there are a variety of complementary explanations, which lead EMW to conjecture that "exchange rate models are not as bad as you think":

1. *EW's (2005) important and influential conjecture.* The results in table 6C2.10, panel A, provide additional empirical evidence that exchange rates do have predictive content for fundamentals, which is an important piece of evidence used in EW (2005) in favor of net present-value models. The results in table 6C2.10, panel B, are also consistent with EW's (2005) conjecture—that high discount factors might undermine

the explanatory power of fundamentals so that exchange rates are better forecasted by a random walk.[9]

2. *Parameter instabilities.* The latter results are also compatible with the empirical evidence discussed in table 6C2.8, that fundamentals may have predictive content for future exchange rates, but this relationship is time-varying and therefore very difficult to exploit for forecasting purposes. The possibility that time variation was empirically relevant in predicting exchange rates was first explored in Meese and Rogoff (1988). We have shown that parameter instabilities are important and might provide another piece of evidence that demonstrates why economic models of exchange rate determination do not work so well in forecasting.[10]

3. *Parameter estimation error.* As discussed above, when comparing a model estimated with a serial correlation correction for the cointegrating vector with a random walk, the random walk's forecasts could be better because the bias in the estimate of the serial correlation could more than offset the gain in exploiting the information on the fundamentals—if the fundamentals are not too informative regarding the exchange rate fluctuations. Therefore, the random walk model might forecast better even if the model is true (see Rossi 2005b).

- *To summarize, although the Meese and Rogoff's stylized fact is still present in the data, we have several possible conjectures for it.*

5 Conclusions

Even though the Meese-Rogoff stylized fact is still alive, we now have a variety of conjectures regarding why it might be the case that economic models do not outperform a random walk in forecasting exchange rates: (a) high discount factors might undermine the predictive content of fundamentals for exchange rates, even theoretically (EW's [2005] important result); (b) widespread instabilities in predictive regressions make it hard, anyway, to use such predictive relationships to forecast exchange rates out of sample; (c) parameter estimation error (Rossi 2005b).

This note showed that, while it is unclear how much variability of exchange rate fundamentals can be explained by looking at the variance ratios, the evidence on Granger-causality tests is quite robust to both uncertainty in the estimation and parameter instabilities. However, there is some evidence that fundamentals have a predictive but time-varying content for future exchange rates.

Does this imply, as in EMW's (2007) words, that: "beating a random walk in forecasting is too strong a criterion for accepting an exchange rate model"? It is certainly a strong criterion, because, as shown by EW (2005) and Rossi (2005b), there are reasons why a random walk may forecast better than the economic model, even if the economic model is true. Nevertheless, out-of-sample forecast comparisons seem to be an important reality check to evaluate the performance of models, and are becoming a standard for evaluating the empirical performance of macroeconomic models as well (Del Negro, Schorfheide, Smets, and Wouters 2007). Understanding and overcoming the reasons of such poor forecasting performance seem to be important tasks for future research, and could be addressed in a variety of ways—for example, by using high-frequency data to capture short-lived effects of news, or time-varying techniques to exploit the time variation in such predictive relationships.

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Endnotes

1. Variance ratios can be longer than 1 in the presence of unmodeled fundamentals (see EW 2004).

2. Throughout, results are similar if the Granger-causality tests are not performed on the constant.

3. Stock and Watson (2003) similarly document widespread instabilities in macroeconomic data.

4. To partially address this issue, Engel and West (2005) report subsample Granger-causality tests, but such results are conditional on knowing the time of the structural break, which in principle is unknown.

5. The concern is related to similar findings in univariate time series models: Corradi, Swanson, and Olivetti (2001) show that out-of-sample tests of equal predictive ability are valid in the presence of cointegration, but Rossi (2005b) shows that they can be misleading when cointegration does not hold. Although there are no similar analyses undertaken in the context of panel regressions, we expect similar results to hold.

6. The fact that the Clark and West (2006) statistic will turn out to be oversized in the presence of roots equal to 1 is not special to their statistic nor a drawback of their procedure: Clark and West (2006, 161) clarify that their results hold if the variables are covariance stationary. Similar results are expected to hold for any tests of out-of-sample forecast comparison. Here we implemented the Clark and West (2006) statistic with an HAC correction for serial correlation with a bandwidth equal to $(h - 1)$ to take into account the fact that h-step-ahead forecasts have a moving average serial correlation structure of order $(h - 1)$.

7. The table reports the Diebold and Mariano (1995) and West (1996) statistic for forecast error comparisons without the adjustment for parameter estimation error. The reason is that the models are nested under the null, so the test statistic is not informative except qualitatively for the fact that when it is positive it means that random walk forecasts better. The critical values used to assess significance are from Clark and McCracken (2001).

8. I thank Ken Rogoff for pointing this out to me.

9. Results are robust to not including the lagged dependent variable.

10. Rossi (2006) examined forecast-combination techniques and models with time-varying parameters and found some evidence that exploiting parameter instability might help in forecasting exchange rates.

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Discussion

The comments during the discussion on "Exchange Rate Models Are Not as Bad as You Think," by Charles Engel, Nelson Mark, and Kenneth West, generally questioned which variables were included as fundamentals and the interpretation of the authors' tests.

Olivier Blanchard, noting that there are large current account surpluses and deficits around the world, posited that the terms of trade of these countries would have to adjust. This adjustment mechanism, relating the current account and the exchange rate, is much discussed in the open-economy macroeconomic literature. However, this paper, Blanchard noted, did not include the current account as a fundamental. Indeed, he said, the paper finds that purchasing power parity, which basically ignores this link, "works quite well." Blanchard wondered why this disconnect exists between the literature and this paper's empirical tests.

Responding to Blanchard's comments, Engel disputed the idea that the U.S. exchange rate will depreciate much further due to the current account imbalance, citing his earlier work. As a result of this belief, the authors did not include the current account as an independent variable. Kenneth Rogoff, however, mentioned his own previous paper, which had added the current account to the set of fundamentals. This variable, along with purchasing power parity, consistently was shown to have a significant impact on exchange rates. He asserted that this result held across dozens of papers.

Daron Acemoglu focused on the interpretation of the authors' Granger causality tests. The exchange rate, under some weak assumptions, would follow a process resembling a random walk. In turn, the authors used the exchange rate to forecast future fundamentals. Acemoglu wondered whether the Granger causality tests could distinguish be-

tween this and the possibility that the exchange rate directly affected the future fundamentals. The latter interpretation made sense because many of the fundamentals were dependent on monetary policy, like interest rates and output.

Kristin Forbes, observing that announcements of macroeconomic data are said to impact the exchange rate, questioned whether policymakers' announcements have been shown to be important. The seeming lack of papers documenting this effect is especially curious, since policymakers believe that their specific statements could have dramatic effects on the economy.

In addition to wanting greater clarification on the specifics of the null test with respect to varying parameters, Julio Rotemberg questioned the importance of one result; namely, the authors' finding that people's expectations of future fundamentals are significantly correlated with the exchange rate. Rotemberg said that it should be obvious that survey data affects the exchange rate. In fact, any financial market will "react to everything," especially given that models with consumer sentiments suggest over-reaction. Kenneth West responded, saying that their results show not just a reaction of exchange rates to changes in people's expectations, but the correct sign as well. However, even though the sign seems accurate, West acknowledged that the magnitude of the coefficient could be off.

Returning to the remarks of the discussants, West made two comments. First, he indicated that the robustness check for parameter instability, proposed by Barbara Rossi, was an interesting idea worth pursuing. He thanked her because, even though they had overlooked the idea, her results seemed to further support their conclusions. Second, West argued about the benefit of Kenneth Rogoff's suggestion to replace the authors' data on people's expectations of future fundamentals with the risk premium. As this term is, at best, poorly measured in the data, the authors' use of survey data to incorporate expectations is justifiable.

Charles Engel continued to discuss Rogoff's critique, saying that there are different, but equivalent, ways to rewrite the model, each yielding different levels of performance, judged by out-of-sample fit. Engel suggested that Rogoff's technique resulted in a poor fit, while the authors' paper used the variable that offered the best fit. As a side note, Rogoff acknowledged that the risk premium could not account for the movements in exchange rates, since it was not nearly volatile enough. There needs to be a leveraging effect, coming through the present discounted values.

Mark Gertler remarked that the Great Moderation in the 1980s resulted in a structural break for financial and real data. Gertler questioned whether the exchange rate data exhibits a similar structural break and how this break affects the performance of the authors' model. Later responding to this question, Nelson Mark indicated that the data by and large do not display any clear structural break in 1984. If anything, the impact of inflation on the exchange rate might have been affected by the Great Moderation. Whereas an increase in inflation made the exchange rate depreciate prebreak, postbreak data shows the opposite result. Mark noted that this conforms to the idea that monetary policy had shifted at that time.

Anil Kashyap concluded the comments session by returning to the discussion of announcements. Kashyap noted that in the finance literature, it is believed that public announcements greatly impact financial variables, including bond yields. The same result, he said, should hold for exchange rates. As a result, Kashyap questioned whether anyone has tried to link major exchange rate movements, unexplained by the unusual fundamentals, with the announcement that caused the change, thus reverse engineering the origin of the movements.