Exchange Rate Models are Better than You Think, and Why They Didn't Work in the Old Days

Charles Engel
University of Wisconsin, NBER and CEPR

Steve Pak Yeung Wu
University of California, San Diego

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Abstract

Exchange-rate models fit very well for the U.S. dollar in the 21st century. A "standard" model that includes real interest rates and a measure of expected inflation for the U.S. and the foreign country, the U.S. comprehensive trade balance, and measures of global risk and liquidity demand is well-supported in the data for the U.S. against other G10 currencies. The monetary and non-monetary variables play equally important roles in explaining exchange rate movements. In the 1970s – early 1990s, the fit of the model was poor but the fit (as measured by t- and F-statistics, and R^2s) has increased almost monotonically to the present day. We make the case that it is better monetary policy (inflation targeting) that has led to the improvement, as the scope for self-fulfilling expectations has disappeared. We provide a variety of evidence that links changes in monetary policy to the performance of the exchange-rate model.

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Emails: Engel: cengel@ssc.wisc.edu, Wu: stevepywu@gmail.com

1. Introduction

In the late 1970s and early 1980s, the "asset market" approach to exchange rate determination emerged. Versions of that model that incorporate slow nominal goods price adjustment posit four important determinants of exchange rates:

1. Monetary models (such as Dornbusch (1976) and the empirical work of Frankel (1979)) highlighted the role of real interest rates. An increase in the U.S. real interest rate leads to a stronger dollar, while increases in foreign real interest rates weaken the dollar, *ceteris paribus*.
2. The asset market approach emphasized that exchange rates, like other asset prices, are forward-looking and incorporate expectations of future "fundamental" determinants. Especially, news about the future monetary policy stance is important in explaining changes in exchange rates.[1]
3. The portfolio balance model, as exemplified by Kouri (1976, 1981) and Branson and Henderson (1985), predicts that as a country's external obligations increase, its currency will depreciate.
4. There is slow, or weak, convergence of exchange rates toward long-run relative purchasing power parity.

While these models initially seemed to have some empirical support, the work of Meese and Rogoff (1983) soon dampened enthusiasm. The title of that paper asked "Exchange Rate Models of the Seventies: Do They Fit Out of Sample?", and the answer was a resounding "No." Cheung et al. (2005), took a second look, and asked "Exchange Rate Models of the Nineties: Are Any Fit to Survive?", and again the answer was negative.

Since 2000, there have been important contributions that document the strength of the U.S. dollar during times of global stress. These studies have shown a strong correlation between the value of the dollar and various measures of this stress such as VIX, the corporate bond spread, deviations from covered interest parity, the convenience yield on dollar bonds, etc.[2]

The first part of our paper documents what we believe is a little-recognized phenomenon: that in the 21st century, a conventional exchange rate model explains the data well. Not only do

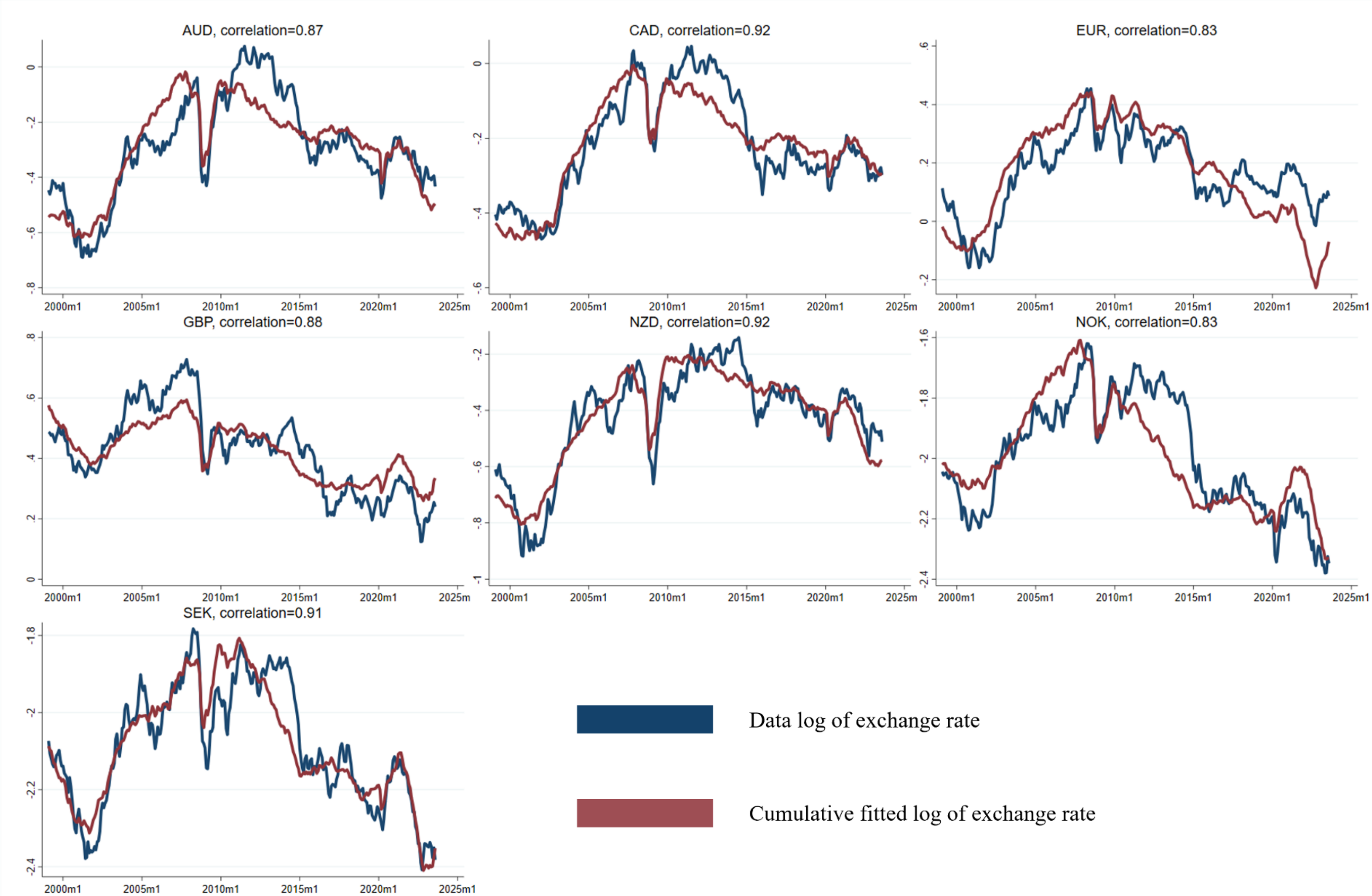
[1] See, for example, Frenkel (1981), and Engel and Frankel (1984).
[2] See, for example, Miranda-Agrippino and Rey (2020), Lilley et al. (2022), Obstfeld and Zhou (2023), Du et al. (2018), Avdjiev et al., Jiang et al. (2021a), Engel and Wu (2023).

measures of global risk and liquidity help account for dollar exchange rates relative to high-income advanced countries, but, controlling for the global risk measures, the traditional variables do as well. In this period, all the variables, both for the U.S. and the "foreign" country (the euro area, the U.K., Canada, Sweden, Norway, Australia, and New Zealand) are highly statistically significant, and the model fits well.

Figure 1 demonstrates how well the model can track the exchange rate dynamics observed in the data. We generate the fitted value of the change in the log of the exchange rate from the model described in Section 2, and cumulatively sum the predicted change in the exchange rate to construct a series of fitted log levels of the exchange rate for each currency. In each subfigure, the blue line represents the empirical log of the exchange rate, and the red line represents the cumulative fitted level. Evidently, the fitted series captures the exchange rate dynamics well. Taking the EUR exchange rate as an example, the fitted values reproduce well the initial appreciation of the U.S. dollar from 1999-2000, followed by the depreciation of the U.S. dollar from 2001 to 2008. The fitted series also matches the sharp appreciation of the U.S. dollar in 2008, 2010, and 2013. Both the data and the fitted series exhibit an appreciation of the dollar from 2013 onwards. The model-implied series also fits the pattern post-2020 very well, mimicking the V-shape from 2021 to 2023. The close correspondence between the red line and the blue lines holds for all other currencies in different sub-periods between 1999 and 2023. The title of each subfigure reports the sample correlation of the two series, indicating that the correlations range from 0.83 for EUR to 0.92 for CAD and NZD.

We emphasize that the model is estimated on monthly changes in log exchange rates. The strong fit does not arise from using longer horizons (such as quarterly changes), which might smooth out unexplained shorter-term movements. We especially note that the model is not fit on levels, where high correlation might arise by coincidence when exchange rates and the explanatory variables are highly persistent. Monthly changes are precisely the horizon over which Meese and Rogoff, and Cheung et al. found previously that models did not fit. We find during this time period that for the dollar relative to the euro, the pound, the Swedish krona, Norwegian krone, and New Zealand dollar, the R^2 is in the range of 0.25 to 0.30, and for the Canadian dollar and Australian dollar, the R^2 is around 0.40 for monthly changes.

Figure 1 Comparing data and model implied exchange rates



Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the data series. Correlations of the two series are reported in the subtitles. The exchange rate is defined as the U.S. dollar price of a foreign currency: an increase in value is a U.S. dollar depreciation.

But the model did not fit over earlier samples. We document this, first by estimating the model over 20-year rolling samples beginning in 1973. In the earlier samples, the fit was poor – the variables are usually statistically insignificant; sometimes when they are significant, they have the wrong sign; and the R^2 values are low. F-tests of the joint significance of the explanatory variables fail to reject the null. But there is a near-monotonic increase in the F-statistics and R^2s as the samples progress in time, and these statistics essentially reach their maximum in the final 20-year sample.

Moreover, we conduct Meese-Rogoff style tests, initially fitting the model over 20-year samples, updating the estimate each month in a rolling sample, then using the Clark-West (2007) statistic to assess the fit continuously. As with the in-sample fit, we find the models do not fit out of sample early on (in the "old days"), but the Clark-West statistic strongly rejects the null of no explanatory power relative to the random walk in later periods.

We treat the dollar/Japanese yen and dollar/Swiss franc exchange rates separately. These two currencies are different, we argue, because of three features that are common to Japan and Switzerland, but not to the other countries: both experienced prolonged periods of deflation; both have engaged in massive sterilized intervention; and, both are considered, along with the U.S., to be "safe haven" currencies. These idiosyncrasies are not well-handled by the standard models, though we argue that our empirical findings point in the direction that these three traits imply for movements of the exchange rates for the dollar against the two currencies.

What accounts for the poor fit of the models in the earlier period, and the excellent fit now? We argue that a change in monetary regime may explain this. Specifically, before countries either explicitly or implicitly adopted inflation targeting, the Taylor principle was not satisfied. When this stability condition is not met, it is possible for self-fulfilling expectations to play a role in real and nominal exchange rate determination, as we demonstrate. We contend that as credibility increased, this phenomenon decreased, and the fit of the standard model improved. This argument is related to three strands of the literature. In studies of the U.S., the "Great Moderation" has been attributed to adoption of policies that resemble inflation targeting beginning in the early 1980s. Clarida et al. (2000) and Coibion and Gorodnichenko (2011), for example, make the case that the Taylor principle did not hold prior to the Volcker era, and, maintain in the context of a closed-economy model, that the greater volatility of the U.S. economy prior to the 1990s was in part due to monetary policy. Most of the other high-income economies did not adopt inflation targeting until later than the Volcker period, so the indeterminacy of exchange rates should persist into the 1990s by this line of reasoning.

The second strand of literature is the "scapegoat" model of exchange rates, originated by Bacchetta and van Wincoop (2004). In that model, markets pay attention to different variables over time, so that the importance of different drivers of the exchange rate are not stable. The scapegoat in the Bacchetta-van Wincoop framework is analogous to the "sunspot" or self-fulfilling expectation channel in an open-economy model in which the Taylor principle is not satisfied. Indeed, the sunspot that influence expectations could be one of the "fundamental" variables that the traditional models posit should drive the exchange rate.

The third strand of the literature regards the declining volatility of exchange rates in floating-rate countries. Ilzetski et al. (2022) and Stavrakeva and Tang (2023), for example, document that dollar exchange rate volatility has gradually declined since the early 1980s. Both studies attribute the decrease in volatility in part to the adoption of inflation targeting, but do not offer a formal explanation of the link. We maintain that stability of exchange rates has increased as the influence of self-fulfilling expectations has declined.

We show that the poor fit of the model in the early part of the sample arise from the statistical insignificance, and sometimes the incorrect sign, of variables associated with monetary policy. While the U.S. real interest rate changes were mildly useful in accounting for exchange rate changes in the first half of the sample, other variables were not. The foreign country real interest rates usually did not contribute in the way models would predict. Notably, expected inflation both in the U.S. and the other countries did not lead to exchange rate changes in the direction that would occur under credible inflation-targeting regimes – that high expected inflation in a country would drive an *appreciation* of that country's currency if its monetary policy was credible. Conversely, the good fit of the model displayed in Figure 1 for the 1999-2023 period can be ascribed in large part to these variables. That is, the good fit does not arise simply from the role of the risk and liquidity variables over this period.

We bolster this analysis by directly examining the monetary policy rules of the U.S. and the other countries in our sample. First, we estimate Taylor rules for each country, but allow for a change in regime. Our method is to use a switching-model with a one-time endogenously determined switch point. We find that for all countries, the early part of the sample corresponds to a period in which the Taylor principle is not satisfied, and the later period to one where it is. The switching date for the U.S. generally comes earlier than for the other countries, consistent with the notion that the coming of the Volcker era marked a change in monetary regime. The improved fit of the model tracks closely to the dates at which regimes switched, as we explain.

We also examine how exchange rates change at the time of announcements of inflation in each country. Here, we measure the unexpected component of the inflation announcement as the difference between the announced inflation result and the median of a survey of traders' expectation of the announcement. When the monetary policy is credible and agents expect the policymaker to target inflation, the currency should appreciate when the announced inflation is greater than expected (as in Anderson et. al., 2003, and Clarida and Waldman, 2008) Since the survey data for the non-U.S. countries only begin in 1997, we cannot track the changes that correspond directly to the switch toward inflation targeting. But we do see that, as a group, the credibility of the monetary rules for the non-U.S. countries appears to have significantly increased relative to findings in previous related studies and between the first half of the sample and the second half, as the reaction of the exchange rate to unexpected inflation is stronger in the second half.

Our study is related to but distinct from the phenomenon of "exchange-rate disconnect…of which the Meese- Rogoff (1983) forecasting puzzle and the Baxter-Stockman (1989) neutrality-of-exchange-rate-regime puzzle are manifestations," to quote Obstfeld and Rogoff (2000), who coined the term. That quote points to two different issues. In the first place, as we have noted, the models that are supposed to explain exchange rates did not work. Second, exchange rate changes seem to have little impact on economic variables. Baxter and Stockman (1989) show that most macroeconomic aggregates did not behave much differently during the period of fixed exchange rates than under volatile floating rates. This suggests that even as exchange rates began to fluctuate, they did not influence GDP, employment, the trade balance, etc., in the way predicted by standard macroeconomic models. Our paper addresses the first issue – we make the case that exchange rates previously could not be well explained by models, and now they can be. We do not address the second leg of "disconnect" – why macro variables, especially in the U.S., do not respond much to exchange rate movements. Other studies offer explanations for this, such as the fact that trade is a small component of the U.S. economy, and that final goods prices are not very responsive to exchange rate movements in the short run.[3]

[3] See Kollmann (2001), Jeanne and Rose (2002), Devereux and Engel (2002), and Itskhoki and Mukhin (2021a,b).

In fact, the second side of disconnect helps to bolster the case for OLS estimation of the exchange rate model. We would like to be confident that the statistical relationship in our regressions between the proxies for the economic fundamentals (monetary policy, debt accumulation, global financial stress) and the exchange rate uncover causality running from the fundamentals to the exchange rate, and not vice-versa. If the exchange rate has only a small causal impact on, for example, inflation or the trade balance, then it is more plausible that the empirical correlation reveals causality from these variables to the exchange rate rather than vice-versa. (In the case of the trade balance, reverse causality is also unlikely because the sign of the estimated coefficient is opposite of the sign that reverse causality would predict.)

Literature Review

Our empirical model ties into five broad strands of the literature on exchange-rate determination: (1) monetary policy; (2) news, especially regarding future inflation or monetary policy; (3) portfolio balance; (4) the role of the dollar in times of global uncertainty; and (6) reversion to purchasing power parity. Since the literature in these areas is, obviously, extensive, the review here is very selective. More complete reviews are available in *Handbook* papers by Maggiori (2022), Du and Schreger (2022), Engel (2014), Gourinchas and Rey (2014), Frankel and Rose (1995), Froot and Rogoff (1995), Dornbusch and Giovannini (1990), Branson and Henderson (1985), Frenkel and Mussa (1985), and Obstfeld and Stockman (1985).

Dornbusch's (1976) classic presentation of the overshooting model is, perhaps, the modern genesis of studies of exchange rates. That paper emphasized the role of contemporary monetary policy changes in driving exchange rates. Frankel (1979) provides empirical support, demonstrating that exchange rates are driven largely by real interest rates. However, Meese and Rogoff (1983) cast doubt on the usefulness of the monetary model in accounting for exchange rate movements, finding that the empirical model of Frankel and related models did not fit the data well out-of-sample. Some recent papers have found direct evidence for the New Keynesian monetary model, such as Schmitt-Grohé and Uribe (2022).

In forward-looking models, exchange rates are connected not just to current economic fundamentals, but, as Frenkel (1981) emphasizes, to news about the future. It is difficult to test the role of the news because there is not a simple measure of news that impacts exchange rates. Some studies, such as Engel and Frankel (1984), Anderson et. al. (2003), and Clarida and Waldman

(2008), examine how announcements of measures of economic aggregates affect the exchange rate. Alternatively, Campbell and Clarida (1987), Engel and West (2005), and Chahrour et al. (2022) investigate whether exchange rates are useful in forecasting other economic aggregates, since news about the future economy may help determine exchange rate movements. Other studies, such as Stavrakeva and Tang (2024) use link expectations from surveys of traders to exchange-rate movements.

The portfolio balance model predicts as a country increases its debt, its currency depreciates. As investors are required to hold more of a country's debt in their portfolio of risky assets, they require a higher expected return. In equilibrium, the currency initially depreciates in response to an increase in debt in order to generate an expected appreciation. This prediction arises in the portfolio balance models of Kouri (1976, 1981), Dooley and Isard (1982), and others that are surveyed by Branson and Henerson (1985). Prominent recent revivals of the portfolio balance model include Gabaix and Maggiori (2015), Della Corte et al. (2016), Greenwood et al. (2023), Gourinchas et al. (2022) and Kremens et al (2023). Also related is the approach of Gourinchas and Rey (2007a,b) that links current account imbalances to future adjustment of the currency price.

The special role of the U.S. dollar during times of global financial stress has been the focus of a very large recent literature. The central role of the U.S., and its status as a safe haven currency, has been explored by, among others, Maggiori (2017), Farhi and Maggiori (2018), Rey (2015, 2016), Gourinchas and Rey (2022), and Kekre and Lenel (2021). Lilley et al. (2023) and Obstfeld and Zhou (2023) find strong empirical support for the relationship between the value of the dollar and measures of financial fragility. Relatedly, several recent papers have found a link between liquidity returns, or convenience yields, and exchange rates, including Du et al. (2018), Krishnamurthy and Lustig (2019), Jiang et al. (2020, 2021, 2024), and Engel and Wu (2023).

There is a very large literature on the adjustment of the exchange rate in the presence of deviations from purchasing power parity. This includes, for example, Rogoff (1996), Taylor (2002), and Taylor and Taylor (2002). The literature concludes that there is evidence of weak convergence of real exchange rates.

Obstfeld and Rogoff (2000) coined the term "exchange-rate disconnect." Models of why the exchange rate is not related to economic fundamentals, and instead may be driven by noise trading have been advanced by Jeanne and Rose (2002), Devereux and Engel (2002), and Itskhoki and Mukhin (2021a,b).

Rose (2007, 2014) makes the case that inflation targeting by central banks around the world has evolved into a new, decentralized, but stable exchange rate system.

Taylor (1999) and Clarida et al. (2000) have argued that the Taylor principle for monetary policy in the U.S. was not satisfied prior to the Volcker era. The "Great Moderation", according to this reasoning, can be explained in part by the Fed's adherence to a stable inflation targeting regime beginning in the early 1980s. Studies that have considered this avenue include Orphanides (2004), Coibion and Gorodnichenko (2011), Bhattarai et al. (2016), and Hirose et al. (2020).

In contemporaneous work, Kekre and Lenel (2024b) build a general equilibrium model that reproduces many of the features of the basic empirical model we estimate here. In particular, the study emphasizes the role of real interest rates as a driver of real exchange rates. The paper does not explicitly incorporate sticky nominal prices, and hence there is no role for monetary policy, though it is not incompatible with a New Keynesian extension of the model, as in Kekre and Lenel (2024a). The study does not attempt to explain why exchange-rate models fit poorly in the 1970s-1990s.

We estimate a quite standard model for exchange-rate determination, augmenting the fundamentals introduced in the 1970s and 1980s with the global risk and liquidity variables the more recent literature has emphasized. In section 2, we recount how each of these variables should affect dollar exchange rates, and then describe our empirical proxies for the fundamentals. Section 3 reports our findings for the model: the fit of the model for the years since the advent of the euro in January 1999 and rolling regressions over 20-year periods starting in 1973. This latter exercise displays the overall fit of the model as measured by R^2 and F-statistics, and of the statistical significance of individual explanatory variables. A subsection considers extensions, including the special cases of the yen and Swiss franc. In section 4, we first review how the failure of monetary policy to satisfy the Taylor Principle may lead to self-fulfilling expectations, and then we present evidence that supports the claim that the improved fit of the models coincides with stricter inflation targeting.

2. An exchange rate model that works

In this section, we present our empirical setup on the exchange rate of the U.S. dollar against the other "G10" currencies: Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), the euro (EUR), U.K. pound sterling (GBP), Japanese yen (JPY), Norwegian krone (NOK), New Zealand dollar (NZD), and Swedish krona (SEK). We start by showing a regression model that works well in explaining the U.S. exchange rate post-1999. We then show that the relationship is not as strong in the past, both in terms of in-sample statistics and out-of-sample fit as in Meese and Rogoff (1983).

a. The empirical exchange rate model

To investigate the empirical relationship of the exchange rate with macro fundamentals, we estimate the following monthly regression from January 1999 to August 2023:

$$(1)\quad \Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta \eta_t + u_t$$

Δ is the first difference operator and represents a one-month difference. The log of the nominal exchange rate, denoted as s_t, is the U.S. dollar price of a foreign currency. An increase in s_t is a depreciation of the U.S. dollar. The real exchange rate is defined as $q_t = s_t + p_t^* - p_t$ where p_t and p_t^* are the U.S. and foreign consumer price indexes. Inflation over the previous 12 months in the U.S. and foreign country is defined as $\pi_t = p_t - p_{t-12}$ and $\pi_t^* = p_t^* - p_{t-12}^*$. The real interest rates are defined as $r_t = i_t - \pi_t$ and $r_t^* = i_t^* - \pi_t^*$ where i_t and i_t^* are the 3-month U.S. and foreign government bond interest rates. To capture the relationship of global risk aversion, we extract the first principal component from five risk measures, including Gilchrist and Zakrajšek (2012) spreads, Moody's Aaa and Baa corporate bond minus Fed Fund rate spreads and Moody's Aaa and Baa corporate bond minus 10 Year Treasury. TB/GDP is the U.S. trade balance (which includes both goods and services) divided by GDP. η_t is a measure of the convenience yield on U.S. Treasury 1-year bonds relative to the convenience yield in the foreign country, measured as the 1-year Treasury basis as in Du et al. (2018) and Engel and Wu (2023).

The model we rely on draws directly from the empirical literature on exchange rates, which in turn draws on fully-specified open-economy macro models. We will not rederive these models, instead describe heuristically how each variable affects exchange rates.

1. *Real interest rates*. An increase in the home real interest rate leads to an appreciation of the home currency. Conversely, an increase in the foreign real interest rate leads to a depreciation of the home currency. Since the dependent variable in our regression is the change in the log of the exchange rate, we include the change in the U.S. real interest rate and the change in the foreign real interest rate as independent variables.

   We need a measure of expected inflation in the U.S. and all the foreign countries to construct the real interest rates. We use simply the inflation rate over the previous year as our measure. This has the benefit of giving us a consistent and easily reproducible measure for all countries.

   While we associate real interest rates with monetary policy decisions, we do not interpret their movement as (necessarily) representing shocks to monetary policy. Rather, these changes are likely to arise from monetary policy responses to macroeconomic conditions.

2. *Inflation*. The current stance of monetary policy, as measured by the real interest rate, is not the only way in which monetary policy may affect exchange rates. Expectations of future monetary policy matter as well. We include either inflation over the past year, or the monthly change in inflation over the past year, for the U.S. and the foreign country in the regression. In both cases, an increase in the inflation variable in the U.S. should lead to an appreciation of the dollar, and an increase in the inflation variable in the foreign country should lead to a dollar depreciation – assuming monetary policy is credible.

   Keeping in mind that the dependent variable is the change in the exchange rate, if we include changes in the inflation rate, we are positing that high inflation today (in levels) in the U.S. implies a stronger dollar today (in levels.) Since we already include the real interest rate, the additional kick coming from the inflation expectation variable represents how markets believe future monetary policy will react to high inflation. When the central bank credibly targets inflation, higher inflation today implies higher future real interest rates, which in turn strengthens the dollar today.

In the alternative specification, we include levels of inflation, rather than changes, to explain changes in the exchange rate. This implies that the price level affects the exchange rate today. If the central bank is a strong inflation targeter, or a price-level targeter, then when prices are high, markets believe future real interest rates will be high, and the currency will be stronger. In other words, if past inflation has been high and price levels are out of line with targets, markets expect future tighter monetary policy.

3. As in the portfolio balance model, higher *external debt* implies a weaker currency. Since the dependent variable is the change in the exchange rate, ideally we would include the change in the external debt as an independent variable. We could measure the accumulation of external debt by the current account deficit, but the current account is not reported monthly. For the U.S., there is a monthly balance of trade in goods and services, which should equal the current account less net international factor payments. This is the variable we use – a higher trade balance should be associated with a decline in the change in the exchange rate.

   Monthly trade balances on goods and services are generally not available for most of the other countries in our sample. If all foreign countries were identical, this would not matter, because the U.S. trade balance would equal the negative of the rest of the world trade balance. In that case, including the foreign trade balance would simply be double counting the effects of changes in the U.S. external debt position. Problems arise when each country deviates from the average of foreign countries. That variable is in the residual of our regression. It could pose a problem, for example, if country X tends to have a high deficit relative to the rest of the world average when the U.S. tends to have a deficit. In that case, the trade balance variable for the U.S. may not successfully contribute to explaining exchange rate movements.

   Also, note that we measure changes in external debt as the accumulation from a broad measure of the trade balance, which does not include valuation changes. This is desirable because it avoids directly "explaining" exchange rates by changes in the value of debt caused by changes in the exchange rate.

4. Measures of *global risk*. Many recent studies have included variants of financial market variables meant to capture global financial stress or uncertainty, such as VIX. The rationale is that the U.S. is a safe haven, so the dollar should appreciate during these times. Various

measures of the corporate bond spread over U.S. Treasuries have proven to be good proxies for this global stress, and these are useful for our purposes because we are able to construct such a measure going back to 1973 when our exchange rate sample starts. Increases in the spread should be associated with an appreciation of the dollar.

5. *Liquidity or convenience yield*. Several recent studies have found that measures of heightened global demand for liquidity are associated with a stronger dollar. We consider two different measures. The first is the "convenience yield" on U.S. Treasury assets. When there is an increase in liquidity demand, markets race to the most liquid asset, U.S. Treasuries. The yield on these bonds fall compared to other interest-earning assets (such as government bonds from foreign countries), so the convenience yield rises and the dollar appreciates.

   Alternatively, we use the liquidity ratio, from Bianchi et al. (2021), to measure the increase in demand for liquidity by financial institutions. This is a measure of liquid dollar reserves and U.S. Treasuries held by commercial banks in the U.S., relative to short-term liabilities. An increase in this ratio is indicative of an increased demand for liquidity and should be associated with an appreciating dollar.

   We note that neither of our measures of liquidity have long time series. We use them in the baseline regressions we report for the post-1999 data, but we do not include them in the rolling regressions of 20-year samples that begin in 1973 in the latter subsection.

6. *Lagged real exchange rate*. We include the lagged real exchange rate as an error correction term. When the real exchange rate is far out of line from its unconditional mean, we posit that some of the adjustment occurs through nominal exchange rate changes. This effect has proven to be weak in many empirical studies of advanced-country exchange rates.

b. Data

Our analyses focus on two sample periods. The first one is January 1999-August 2023. The second one is March 1973-August 2023. We study the U.S. exchange rate verse the rest of the "G10" currencies: Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), the euro (EUR), U.K. pound sterling (GBP), Japanese yen (JPY), Norwegian krone (NOK), New Zealand dollar (NZD), and Swedish krona (SEK). We use monthly average data to smooth out shorter-term movements. We obtain nominal exchange rate data from Federal Reserve Board of Governors

H.10 series, which is sourced from the Federal Reserve Economic Data (FRED). Home and foreign consumer price indexes (CPI) are obtained from the IMF International Financial Statistics. Inflation rates over the previous twelve months are computed as the 12-month log difference of CPI. The nominal government bond interest rate are obtained from Global Financial Database (GFD). To construct the $RISK_t$ variable that captures global risk aversion, we extract the first principal component from five risk measures that goes back to 1973. These include Gilchrist and Zakrajšek (2012) spreads, Moody's Aaa and Baa corporate bond minus Fed Fund rate spreads (FRED series: AAAFF, BAAFF) and Moody's Aaa and Baa corporate bond minus 10 Year Treasury (FRED series: AAA10Y, BAA10Y). We use two measures of liquidity/convenience. First, the liquidity yield measure in Engel and Wu (2023), which is the CIP deviation of 1-year government bond rate. Second, the liquidity ratio in Bianch e al (2021), which is constructed by the sum of U.S. dollar financial commercial paper (FRED series: DTBSPCKFM) and short-term funding to U.S. banks is demand deposits (FRED series: DEMDEPSL) divided by the sum of reserves held at Federal Reserve banks and government securities (Treasury and agency) held by commercial banks (the sum of TOTRESNS and USGSEC from FRED.) The trade balance variable includes both goods and services and the frequency is monthly after 1992 and quarterly before 1992. We interpolate both the trade balance pre-1992 and quarterly nominal GDP variables linearly to obtain monthly observations. Appendix Table 1 provides the detailed data sources and sample periods of each of the variables.

3. Estimated model

a. Model estimated in post-1999 data

Table 1 presents the baseline regression results. Each column displays the regression coefficients for the column-head currency's exchange rate with the U.S. dollar. Column (8) and column (9) are for panel regressions of all seven currencies with and without fixed effects. The R^2 and F-statistics of these regressions are high. The R^2s range from 0.24 for GBP to 0.38 for AUD. The F-statistics are all above 10, which correspond to a p-value smaller than 0.0001 and indicate a very high joint significance of the variables in explaining exchange rates.

The real interest rate coefficients are estimated to be strongly statistically significant with expected signs, except for the real interest rate for NOK. All the U.S. real interest rates are estimated to be significantly negative (at the one percent level), indicating a higher U.S. interest rate is associated with an immediate U.S. dollar appreciation. For example, the coefficient for AUD is -1.12 for the U.S. interest rate and 1.12 for the Australian interest rate, which imply that a one percentage point increase in the annualized U.S. (Australian) interest rate is associated with 1.12% appreciation (depreciation) of the U.S. dollar. It is interesting to note that for most currencies, the absolute value of the coefficient for the U.S. and foreign real interest rates are approximately equal, as we would find in a standard symmetric model.

The inflation coefficients are all estimated with expected signs, except for the foreign inflation of NOK. Importantly, all the U.S. inflation coefficients are negative and significant. This implies that when the U.S. inflation is high, it is associated with an appreciation of the U.S. dollar. For example, the coefficient for AUD is -0.26 for U.S. inflation, which indicates a one percentage point year-on-year inflation is associated with 0.26% appreciation of the U.S. dollar. When a central bank is credible and the main objective of monetary policy is price stability, the current inflation rate is informative about the future path of interest rates. A plausible interpretation of the negative U.S. inflation coefficients is that when the U.S. inflation is high, market participants expect an increase in U.S. interest rate in the near future, resulting in an appreciated U.S. dollar.

The coefficients on risk variable ($RISK_t$) and liquidity/convenience variable (η_t) are estimated to be negatively significant for most cases. A higher value of $RISK_t$ implies an increase in global risk aversion, resulting in a high excess return of investing in foreign currency through an immediate U.S. dollar appreciation. The liquidity/convenience variable, in this case measured as the relative convenience yield on 1-year government bonds, captures a related exchange rate mechanism. When there is a higher demand for convenience assets, such as a high demand for collateral services (Devereux et al 2023), global flight to safety (Jiang et al 2021) or funding uncertainty (Bianchi et al 2021), the U.S. dollar appreciates.

The coefficients on TB/GDP are negatively significant. It indicates that when there is a trade balance improvement for the U.S., the dollar appreciates. This is consistent with the portfolio balance model prediction that when a country's external debt decreases, its currency will strengthen. The coefficient of -0.54 for the panel regression indicate if the TB/GDP improves by one percentage point, the U.S. dollar immediately appreciates by 0.54%.

Finally, we find some evidence of long run PPP holds, which is consistent with the recent findings of Eichenbaum et al (2021). The negative sign implies when the real exchange rate is above its mean, the nominal exchange rate appreciates subsequently to bring the real exchange rate back to the long run mean. The coefficients are estimated to be significant for EUR, GBP, NOK and the panel fixed effect specification.

In Appendix C, Tables C1 and C2, we present a simple metric of the contribution of the monetary variables (real interest rates and inflation) and the non-monetary variables. In those tables, we fit the model to each group of variables separately. The R^2s show that monetary and non-monetary variables roughly contribute equally to the overall fit of the exchange rate. In addition, regression coefficients are estimated with the expected sign and are significant in each regression.

Figure 1, discussed in the introduction, shows the fitted values from these regressions. We noted at the outset the very high correlation between the fitted value in levels (from cumulating the model's fitted values for the change in the log of the exchange rate) and the level of the exchange rate. While there is no deterministic trend in these exchange rates, the correlations in levels (in the neighborhood of 0.90) compared to the R^2 for the model estimated in monthly changes (still an impressive range of 0.24 to 0.38) reflects the persistent swings in the levels.

Tables 2 and 3 present two alternative specifications of the basic model. In Table 2, the measure of global liquidity demand is the liquidity ratio of U.S. commercial banks, from Bianchi et al. (2021). Table 3 retains the original convenience yield measure, but includes monthly changes in inflation, rather than inflation levels. In short, the performance of the models, in terms of overall fit and statistical significance of the individual variables is very similar to what we found in Table 1. The appendix reports the estimates of these models without the liquidity variables, which are not available for most of the post-1973 floating exchange rate era (and, hence, are not used in the following sub-section.) The overall fit of the models is slightly worse with the omission of this variable, but the R^2 values and statistical significance of the remaining included variables remain high.

It is unlikely that the model estimated here is polluted by reverse causality – that is, that the correlations of the exchange rate with the right-hand-side variables arise because the exchange rate change causes the change in the other variables. None of the central banks for these countries use monetary policy explicitly to target exchange rates. So, it is unlikely that the correlation of the exchange rate with real interest rates reflects a response of policy to exchange rate changes. Even if there were reverse causality, the signs of the estimated coefficients are the opposite of what we would expect if central banks were targeting exchange rates. One would expect that when the currency was depreciating, they would raise the interest rate, but we see the opposite partial correlation in this table.

Similarly, it is notable that we find higher inflation is associated with a stronger currency. The direct effect of exchange rate changes on inflation might be expected to go the other way – depreciation leads to higher inflation. Instead, as we would expect if monetary policy were credible, higher inflation over the previous 12 months is associated with an appreciation of the currency.

In any case, as we noted above, exchange rate changes are thought to have only small effects on macro variables such as inflation or the trade balance. In fact, also in the case of the trade balance, if the causality ran from the exchange rate to the trade balance, we would expect a positive relationship – depreciation raises the trade balance, but Table 1 shows the opposite relationship. Finally, it is hard to conceive of an argument in which the change in the dollar exchange rate is the cause of global uncertainty or liquidity demand. As the recent literature has emphasized, the dollar movements are likely a symptom not a cause of the global disruptions.

Table 1: Baseline regression with inflation level, and convenience yield

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta \eta_t + u_t$$

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-1.12***	-1.66***	-2.34***	-1.37***	-0.92***	-1.85***	-1.94***	-1.48***	-1.47***
	(-3.83)	(-6.38)	(-8.11)	(-5.32)	(-2.83)	(-6.37)	(-6.53)	(-7.35)	(-7.30)
Δr_t^*	1.12***	1.16***	2.21***	1.80***	1.06***	0.23	0.80**	0.92***	0.94***
	(4.03)	(4.30)	(5.60)	(4.84)	(2.97)	(1.14)	(2.42)	(5.63)	(5.72)
π_t	-0.25**	-0.21*	-0.69***	-0.33***	-0.45***	-0.21*	-0.58***	-0.34***	-0.33***
	(-2.41)	(-1.66)	(-4.88)	(-2.85)	(-3.50)	(-1.94)	(-4.65)	(-4.32)	(-4.31)
π_t^*	0.03	0.14	0.53***	0.14	0.24*	-0.19	0.24**	0.15**	0.18***
	(0.24)	(0.94)	(4.02)	(1.31)	(1.76)	(-1.48)	(2.57)	(2.13)	(2.83)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01***	-0.01***	-0.02***	-0.02***	-0.02***	-0.02***	-0.02***
	(-10.19)	(-9.60)	(-2.77)	(-4.61)	(-6.67)	(-7.56)	(-5.92)	(-7.70)	(-7.44)
q_{t-1}	-0.01	-0.01	-0.02**	-0.03***	-0.01	-0.03***	-0.01	-0.01**	-0.00
	(-1.34)	(-1.50)	(-2.58)	(-2.72)	(-1.50)	(-2.87)	(-1.33)	(-2.19)	(-0.45)
$\frac{TB}{GDP_t}$	-0.48***	-0.45***	-0.63***	-0.73***	-0.37*	-0.66***	-0.80***	-0.54***	-0.48***
	(-2.76)	(-3.71)	(-3.86)	(-3.49)	(-1.91)	(-3.16)	(-4.00)	(-4.26)	(-3.80)
$\Delta \eta_t$	-1.92**	-2.33***	-0.86	-1.52*	-1.56**	-1.20*	-1.04	-1.38**	-1.45**
	(-2.09)	(-2.93)	(-0.92)	(-1.76)	(-2.11)	(-1.77)	(-1.54)	(-2.24)	(-2.33)
N	296	296	295	296	296	296	296	2071	2071
F	21.80	21.45	13.30	11.56	11.33	16.80	13.12	22.65	21.50
R2	0.38	0.37	0.27	0.24	0.24	0.32	0.27		0.25
R2_adj	0.36	0.36	0.25	0.22	0.22	0.30	0.25		
R2_within								0.25	

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the measure of the U.S. convenience yield relative to the foreign country, using 1-year government bond rates, as in Engel and Wu (2023).

Table 2: Alternative specification of the baseline regression with inflation level, and liquidity ratio

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta Liqudity\ Ratio_t + u_t$$

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-1.05***	-1.62***	-2.17***	-1.39***	-0.74**	-1.80***	-1.81***	-1.36***	-1.36***
	(-3.40)	(-5.78)	(-7.54)	(-5.12)	(-2.22)	(-5.92)	(-5.93)	(-6.90)	(-6.92)
Δr_t^*	0.95***	1.18***	2.23***	1.53***	1.07***	0.20	0.81**	0.85***	0.87***
	(3.13)	(4.06)	(5.67)	(3.85)	(2.91)	(0.99)	(2.32)	(5.02)	(5.09)
π_t	-0.30**	-0.30**	-0.67***	-0.47***	-0.40***	-0.26**	-0.63***	-0.38***	-0.38***
	(-2.59)	(-2.18)	(-4.65)	(-3.42)	(-2.70)	(-2.24)	(-4.72)	(-4.74)	(-4.85)
π_t^*	0.09	0.24	0.48***	0.23*	0.14	-0.15	0.29***	0.17**	0.22***
	(0.59)	(1.41)	(3.61)	(1.90)	(0.86)	(-1.12)	(2.87)	(2.39)	(3.36)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01**	-0.01***	-0.02***	-0.02***	-0.02***	-0.02***	-0.02***
	(-9.36)	(-9.16)	(-1.98)	(-3.86)	(-6.31)	(-6.85)	(-5.60)	(-7.79)	(-7.53)
q_{t-1}	-0.01	-0.01	-0.03***	-0.04***	-0.02**	-0.03***	-0.01	-0.02**	-0.00
	(-1.49)	(-1.15)	(-2.65)	(-2.96)	(-2.04)	(-2.67)	(-1.39)	(-2.51)	(-0.71)
$\frac{TB}{GDP_t}$	-0.50***	-0.52***	-0.60***	-0.93***	-0.38*	-0.68***	-0.87***	-0.57***	-0.50***
	(-2.78)	(-4.01)	(-3.73)	(-3.76)	(-1.94)	(-3.11)	(-4.11)	(-4.50)	(-4.02)
$\Delta LiqRatio_t$	-0.06	-0.03	-0.09**	-0.09**	-0.06	-0.06	-0.07*	-0.07**	-0.07**
	(-1.43)	(-1.21)	(-2.48)	(-2.43)	(-1.28)	(-1.49)	(-1.76)	(-2.35)	(-2.35)
N	271	271	271	271	271	271	271	1897	1897
F	20.13	19.11	13.42	11.59	11.22	15.51	12.74	28.82	26.45
R2	0.38	0.37	0.29	0.26	0.26	0.32	0.28		0.25
R2_adj	0.36	0.35	0.27	0.24	0.23	0.30	0.26		
R2_within								0.26	

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Feb 2001 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. $\mathrm{r_t}$ and $\mathrm{r_t^*}$ are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the liquidity ratio from U.S. commercial banks, as defined in Bianchi, et al. (2021).

Table 3: Alternative specification of the baseline regression with the change of inflation rate, and convenience yield

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \Delta \pi_t + \beta_4 \Delta \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta \eta_t + u_t$$

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-4.11***	-3.91***	-3.47***	-2.43***	-4.35***	-2.61***	-3.37***	-2.89***	-2.88***
	(-5.56)	(-5.96)	(-5.00)	(-3.63)	(-5.03)	(-3.46)	(-4.37)	(-4.57)	(-4.57)
Δr_t^*	4.38***	3.52***	2.85***	3.47***	4.18***	0.07	2.41***	1.36**	1.41**
	(5.56)	(5.00)	(3.45)	(4.58)	(4.91)	(0.23)	(2.75)	(2.16)	(2.20)
$\Delta \pi_t$	-3.62***	-2.66***	-1.76**	-1.57**	-4.33***	-0.78	-2.01**	-1.81***	-1.83***
	(-4.46)	(-3.73)	(-2.24)	(-2.10)	(-4.56)	(-0.96)	(-2.39)	(-2.61)	(-2.66)
$\Delta \pi_t^*$	3.45***	2.74***	0.62	2.01**	3.56***	-0.41	1.59*	0.58	0.64
	(4.25)	(3.81)	(0.64)	(2.27)	(3.99)	(-1.09)	(1.76)	(0.88)	(0.95)
$\Delta RISK_t$	-0.04***	-0.02***	-0.01***	-0.01***	-0.03***	-0.03***	-0.02***	-0.03***	-0.03***
	(-10.14)	(-9.57)	(-3.80)	(-4.71)	(-8.26)	(-7.33)	(-6.55)	(-9.39)	(-9.34)
q_{t-1}	-0.00	-0.01	-0.02*	-0.01	-0.01	-0.01	-0.01	-0.01	0.00
	(-0.57)	(-1.54)	(-1.96)	(-1.27)	(-0.92)	(-1.42)	(-0.57)	(-1.54)	(0.20)
$\frac{TB}{GDP_t}$	-0.30*	-0.34***	-0.38**	-0.30*	-0.15	-0.45**	-0.33*	-0.32***	-0.28**
	(-1.91)	(-3.16)	(-2.49)	(-1.77)	(-0.87)	(-2.53)	(-1.84)	(-2.73)	(-2.43)
$\Delta \eta_t$	-2.16**	-2.53***	-1.26	-1.70**	-1.98***	-1.31*	-0.91	-1.62***	-1.68***
	(-2.42)	(-3.25)	(-1.27)	(-1.97)	(-2.72)	(-1.87)	(-1.29)	(-2.66)	(-2.75)
N	296	296	295	296	296	296	296	2071	2071
F	25.76	23.93	10.35	11.12	13.44	14.77	10.73	24.19	23.36
R2	0.42	0.40	0.22	0.24	0.27	0.29	0.23		0.24
R2_adj	0.40	0.38	0.20	0.22	0.25	0.27	0.21		
R2_within								0.24	

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the measure of the U.S. convenience yield relative to the foreign country, using 1-year government bond rates, as in Engel and Wu (2023).

b. Model estimation from 1973

In the previous subsection, we have shown that the exchange-rate model is successful for the sample from 1999-2023. In this subsection, we extend the sample to 1973 and perform regressions in a 20-year rolling window to see how the in-sample and out-of-sample performance of the model changes over time. We make two adjustments when we extend the sample. First, extending the model back before 1999 requires us to focus on the deutschemark (DEM) pre-1999, spliced with the euro beginning in January 1999. Second, because of the lack of data for the liquidity/convenience measure, we drop the variable for this exercise.

In-sample analysis

We first provide a snapshot of the pre-1999 regression results in Table 4, which reports the same regression as in Table 1 but changing the sample to March 1973 to Dec. 1998. Table 4 shows a drastically different pattern. Most of the regression coefficients are not statistically significant, and many of the coefficient estimated are opposite of the theoretical prediction. For example, three of the foreign interest rate coefficients (CAD, DEM, GBP) are positive. The overall fit is poor - the R^2 is less than 0.1 for most currencies.

In Figure 2, we plot the *t*-statistics for each of the variables of the 20-year rolling window regression. The *x*-axis represents the start date of the rolling window. That is, the start date of March 1973 reports the *t*-statistics from a regression with the sample period from March 1973 to Feb 1993. The last regression ends with a start date of September 2003. In the left panels, we plot the *t*-statistics of the monetary variables: change of home real interest rate (blue line), change of foreign real interest rate (red line), home inflation rate (green line) and foreign inflation rate (orange line). In the right panels, we plot the *t*-statistics of change of global risk measure (black line), the lagged real exchange rate (grey line) and the U.S. trade balance to GDP (golden line).

It is clear from Figure 2 that the model relationships started to be "well behaved" in the second half of the sample. In the latter half, the home real interest rate and home inflation rate are negatively significant, and the foreign real interest rate and foreign inflation are positively significant. The *t*-statistics of the global risk measure, the lagged exchange rate and the trade balance to GDP ratio all shifted to the negative region (as the model implies) in the second half of the sample. However, these patters do not hold in the first half of the sample. For example, the third subfigure that reports the case of DEM, the *t*-statistics for both the German interest rate and inflation are negative and not significant, which is the opposite sign of what a monetary model would predict. In addition, the *t*-statistics of the global risk variable and lagged real exchange rate are positive in the first 10 years of the sample and the trade balance to GDP is not statistically significant for a long period.

In Figure 3, we report the *F*-statistics for the null of no joint explanatory power of the right-hand-side variables and R^2 for the same 20-year rolling regression in Figure 2. Both increase nearly monotonically over the entire period, reaching approximately their maximum values in the last sample. (An exception to the monotonic increase is the U.K. pound – there is a window in which the fit worsens for samples that begin in the 1990s, before increasing again.) Consistent with the regressions report in Table 1, which has roughly 24 years of sample, the *F*-statistics are very statistically significant in the second half of the sample. However, when looking at the rolling window regression that goes back in time to 1970s to 1980s, the *F*-statistics falls dramatically.

Meese Rogoff out-of-sample fit

In this section, we conduct an out-of-sample fit exercise as in Meese Rogoff 1983. For each time *t*, we estimate the regression:

$$\Delta s_t = \alpha_{t-1} + \beta_{1,t-1}\Delta r_t + \beta_{2,t-1}\Delta r_t^* + \beta_{3,t-1}\pi_t + \beta_{4,t-1}\pi_t^*$$
$$+\beta_{5,t-1}\Delta Risk_t + \beta_{6,t-1}q_{t-1} + \beta_{7,t-1}\frac{TB}{GDP_t} + u_t$$

where each of the α and β are estimated using sample from period $t-1$ to period $t-240$ (a 240-month rolling window). In each rolling regression, we record the prediction error u_t and compare it with the prediction error of a random walk, $u_t^{RW} = \Delta s_t$. Formally, we use Clark-West (2007) statistics to test the null that the root mean square prediction error of our model is equal to that of a random walk model. The Clark-West statistics adjust for the fact that our model nests the random walk model if all regression coefficients are set to zero.

Since our sample starts from March 1973, we use the sample from March 1973 to February 1993 to estimate the first regression and generate the first prediction in March 1993. The sample is kept fixed for 240 months so regression coefficients of the second prediction in April 1993 is estimated using data from April 1973 to March 1993 and so on.

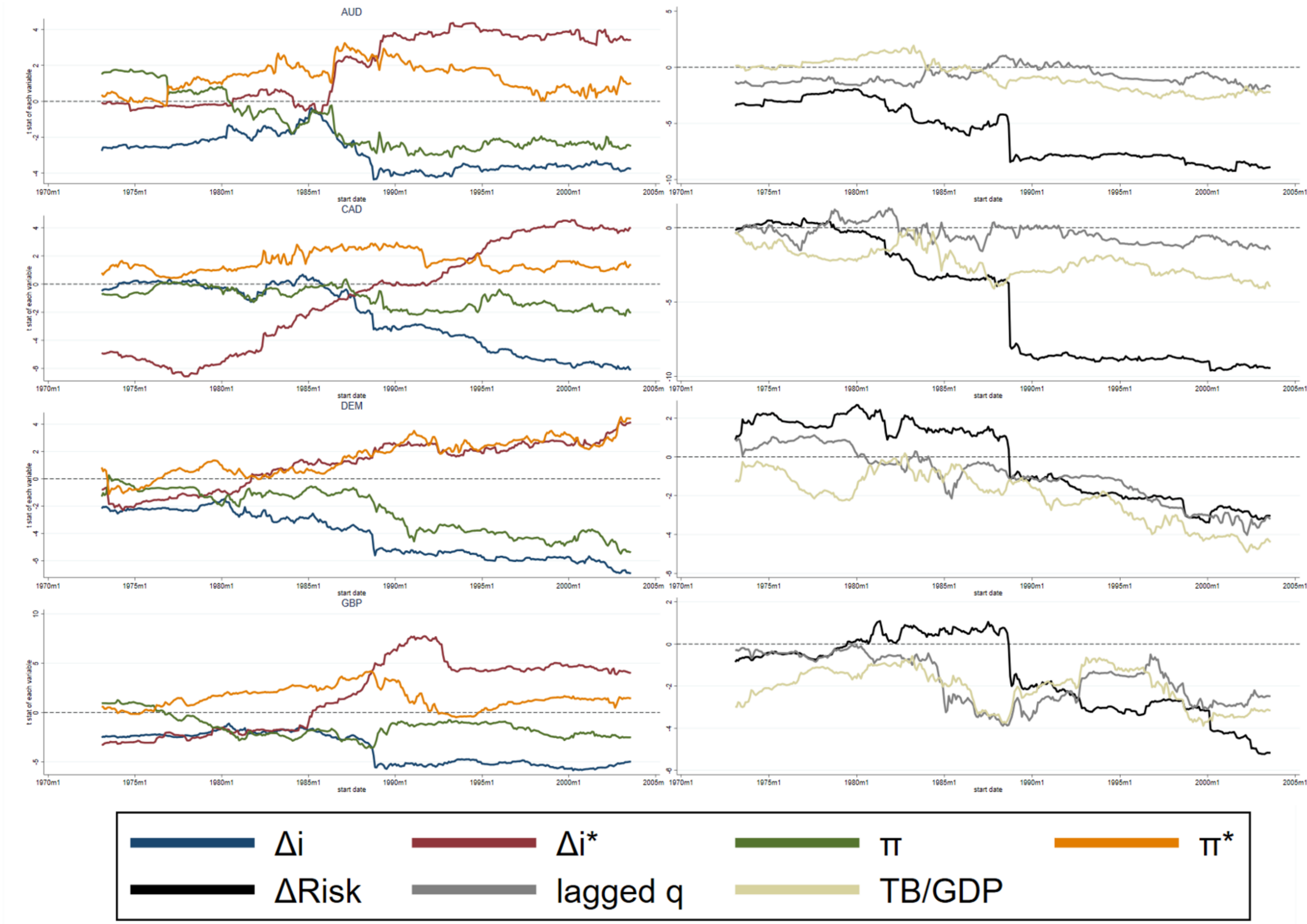
Table 4: Baseline regression with inflation level, sample period Mar 1973-Dec 1998

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + u_t$$

	AUD	CAD	DEM	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-0.72***	-0.03	-0.77**	-0.72**	-0.18	-0.65**	-0.86***	-0.63***	-0.63***
	(-2.84)	(-0.24)	(-2.48)	(-2.59)	(-0.40)	(-2.49)	(-3.17)	(-3.55)	(-3.56)
Δr_t^*	0.03	-0.51***	-0.22	-0.62***	0.11	0.16	0.35***	0.02	0.02
	(0.20)	(-5.80)	(-0.66)	(-3.57)	(0.74)	(0.84)	(2.99)	(0.28)	(0.28)
π_t	0.08	-0.03	-0.09	0.07	-0.20*	0.07	0.02	0.02	0.02
	(1.21)	(-0.76)	(-1.23)	(0.77)	(-1.86)	(1.03)	(0.27)	(0.42)	(0.44)
π_t^*	0.02	0.05	0.26	-0.02	0.12**	-0.06	-0.01	0.01	0.01
	(0.39)	(1.39)	(1.39)	(-0.33)	(2.18)	(-1.08)	(-0.27)	(0.29)	(0.29)
$\Delta RISK_t$	-0.01***	-0.00	0.00	-0.00	0.00	-0.00	-0.00	-0.00	-0.00
	(-3.61)	(-0.30)	(0.93)	(-0.69)	(0.43)	(-0.09)	(-1.09)	(-1.24)	(-1.23)
q_{t-1}	-0.03*	0.00	0.01	-0.01	0.01	-0.01	0.00	-0.00	-0.00
	(-1.78)	(0.11)	(0.56)	(-0.50)	(1.14)	(-0.98)	(0.46)	(-0.20)	(-0.47)
$\frac{TB}{GDP_t}$	0.19	-0.08	-0.77*	-0.53**	0.01	-0.17	-0.42*	-0.26*	-0.26*
	(0.79)	(-0.91)	(-1.82)	(-2.28)	(0.04)	(-0.86)	(-1.92)	(-1.95)	(-1.92)
N	311	311	311	311	205	311	311	1760	1760
F	2.25	6.91	3.52	4.89	1.36	2.08	3.65	3.34	3.45
R2	0.05	0.14	0.08	0.10	0.05	0.05	0.08		0.03
R2_adj	0.03	0.12	0.05	0.08	0.01	0.02	0.06		
R2_within								0.03	

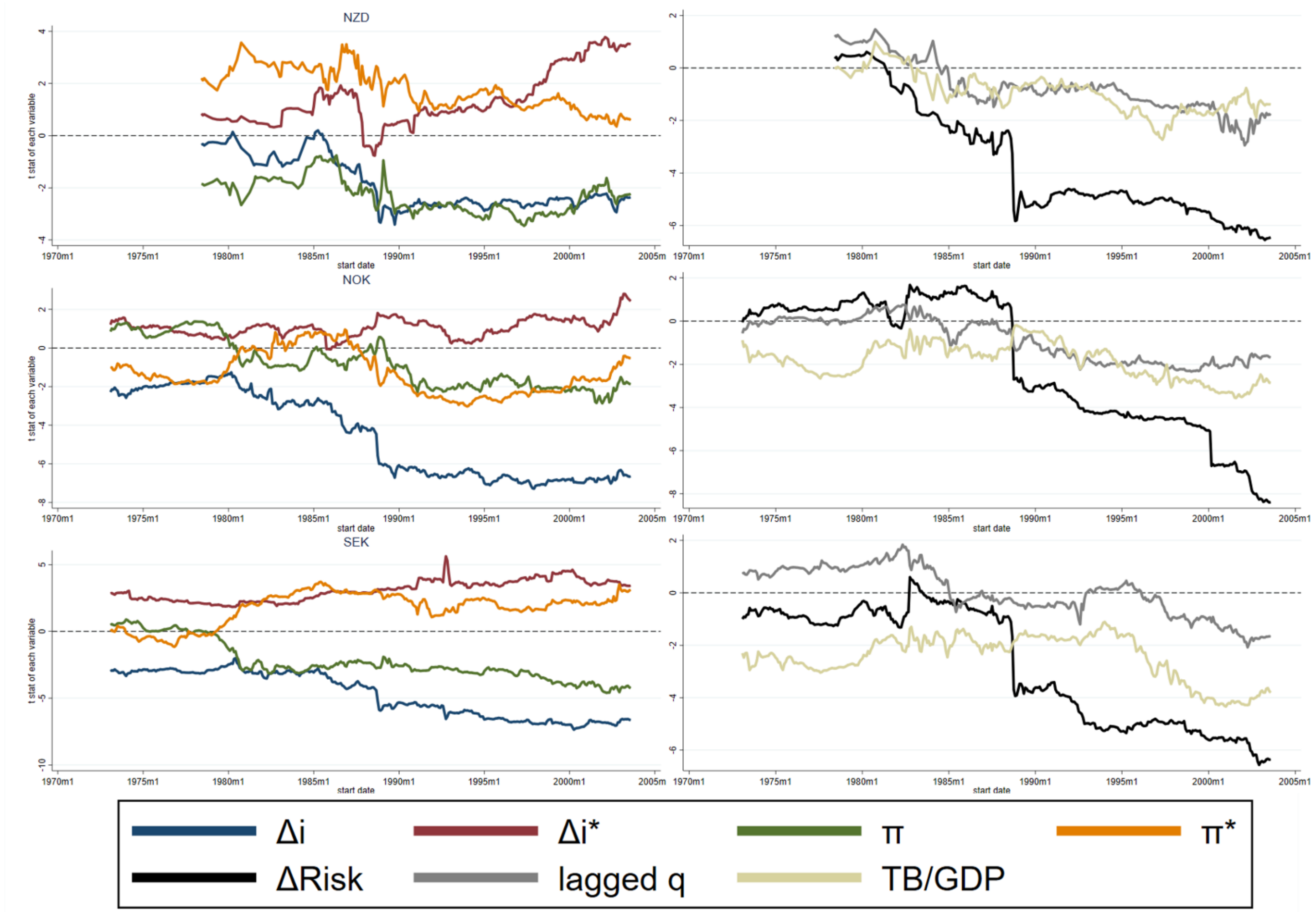
Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Mar 1973 to Dec 1998. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S.

Figure 2: *t* statistics of 20-year rolling window regressions of equation (1)



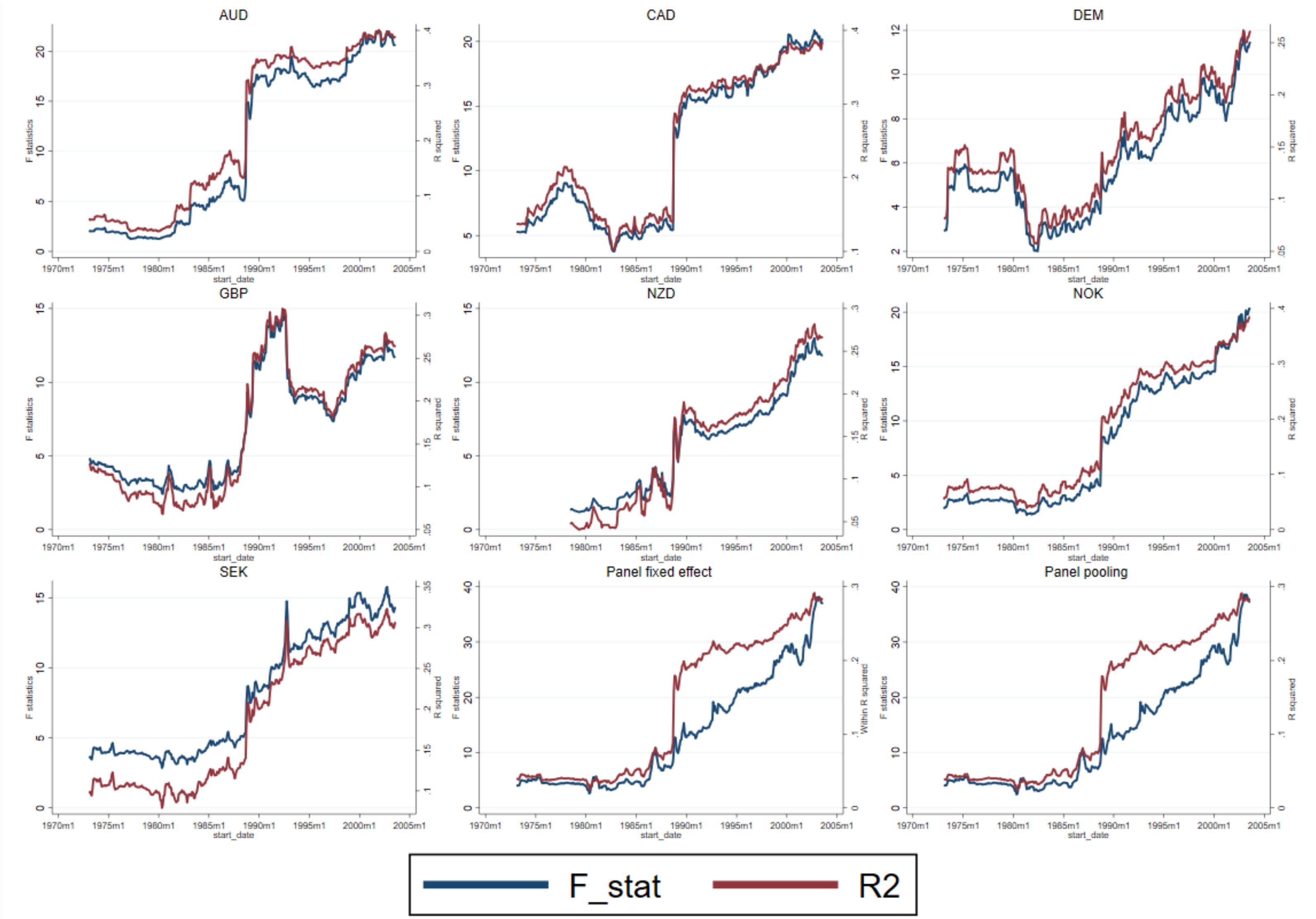
Note: The figure reports the t-statistics of each of the variables in equation (1) with a 20-year rolling window regression. X-axis correspond to the start date of the rolling window regression . Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2003-Aug 2023.

Figure 2: *t* statistics of 20-year rolling window regressions of equation (1), continued



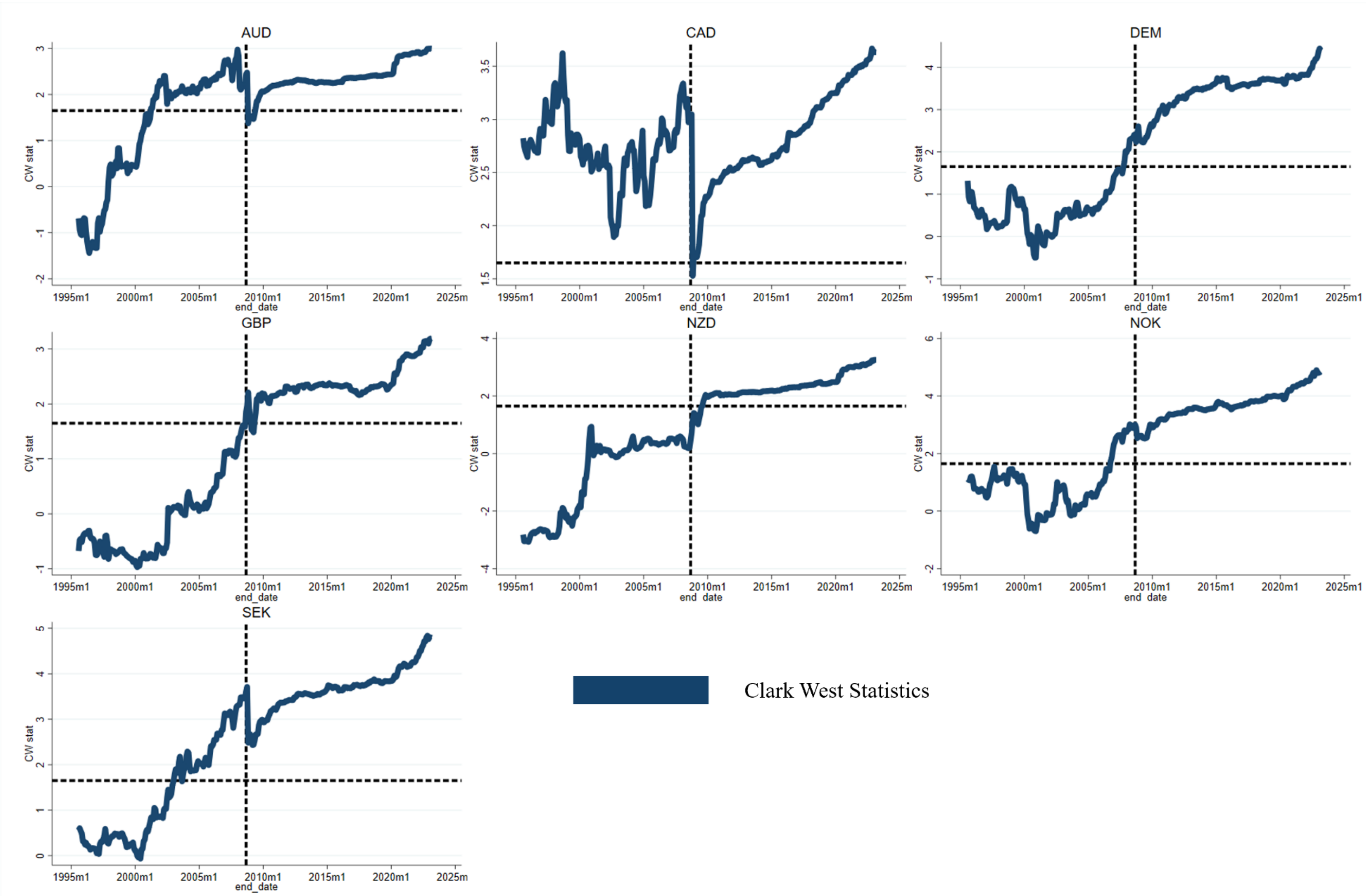
Note: The figure reports the t-statistics of each of the variables in equation (1) with a 20-year rolling window regression. X-axis correspond to the start date of the rolling window regression . Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2003-Aug 2023.

Figure 3: *F*-statistic of 20-year rolling window regressions of equation (1)



Note: The figure reports the F-statistics and R squared in equation (1) with a 20-year rolling window regression. X-axis corresponds to the start date of the rolling window regression. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2003-Aug 2023.

Figure 4: Clark West statistics of out-of-sample fit



Note: The figure reports the Clark-West statistics of each of the variables in equation (1) with 20-year rolling window regressions. X-axis corresponds to sample end period of the Clark-West statistic. The first Clark-West statistic is computed with prediction errors from Mar 1993 to Sep 1995. The sample of prediction errors keep increasing till the end of sample. The last Clark-West statistic is computed with prediction errors from Mar 1993 to Aug 2023.

Figure 4 plots the Clark-West statistics over time. Under the null hypothesis, the Clark- West statistics follow a normal distribution. The horizontal black dash lines indicate the value of 1.65, which is the critical value where the exchange rate model outperforms the random walk at a 5% significance level.

We begin computing root mean square prediction errors and the Clark-West statistics as soon as we have 30 predictions, which occurs in September 1995. We recalculate the statistics and include the additional prediction errors as time progresses. In other words, we continuously evaluate our model from September 1995, October 1995, November 1995, and so on in real-time, assessing if our model generates statistically significantly better predictions than a random walk until the end of the sample. The Clark-West statistics at the end of the sample provide inference about whether the exchange rate model outperforms the random walk model throughout the entire sample.

Consistent with the literature, we find limited evidence that the exchange rate model can beat the random walk in the very beginning of the sample. The Clark-West statistic is significantly positive for the Canadian dollar exchange rate from the beginning of the sample. Except for CAD, the exchange rate model does not outperform the random walk significantly. In fact, there are three currencies — AUD, GBP, and NZD — for which the Clark West statistics are negative, meaning that the root mean square prediction errors of the exchange rate model are worse than the random walk model.

As time goes on, we find that the performance of the exchange rate model gradually increases. For all currencies, the exchange rate model beats a random walk model by the end of the sample significantly at the 1% level (Clark-West statistic above 1.96).

The exact time that the exchange rate model significantly outperforms the random walk model varies across countries. The vertical dashed lines indicate the month of September 2008 when Lehman Brothers collapsed. The figure shows that except for GBP and NZD, the exchange rate model significantly fit better than the random walk model out of sample before 2008, as early as 2003, such as in the case of AUD and SEK.

Table 5: Baseline regression with inflation level, and convenience yield for Swiss franc and Japanese Yen

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta \eta_t + u_t$$

	CHF	JPY		CHF	JPY
Δr_t	-1.42***	-0.36	Δr_t	-2.66***	-2.29***
	(-4.39)	(-1.19)		(-3.56)	(-2.98)
Δr_t^*	0.86*	-0.04	Δr_t^*	-0.14	7.07**
	(1.75)	(-0.10)		(-0.16)	(2.21)
π_t	-0.27**	-0.29***	$\Delta\pi_t$	-1.50*	-2.25***
	(-2.10)	(-2.91)		(-1.76)	(-2.73)
π_t^*	0.35	-0.05	$\Delta\pi_t^*$	-1.13	7.00**
	(1.49)	(-0.30)		(-1.07)	(2.19)
$\Delta RISK_t$	-0.00	0.01***	$\Delta RISK_t$	-0.01	0.01**
	(-0.62)	(3.77)		(-1.63)	(1.99)
q_{t-1}	-0.01	-0.01	q_{t-1}	-0.01	0.00
	(-0.85)	(-0.92)		(-1.24)	(0.12)
$\frac{TB}{GDP_t}$	-0.18	-0.32*	$\frac{TB}{GDP_t}$	-0.19	-0.04
	(-1.05)	(-1.71)		(-1.17)	(-0.26)
$\Delta\eta_t$	-2.40***	-1.48	$\Delta\eta_t$	-2.31***	-2.07*
	(-2.92)	(-1.38)		(-2.76)	(-1.92)
N	296	296	N	296	296
F	4.32	3.17	F	4.72	3.27
R2	0.11	0.08	R2	0.12	0.08
R2_adj	0.08	0.06	R2_adj	0.09	0.06

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. $\Delta \mathrm{r_t}$ and $\Delta \mathrm{r}_t^*$ are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S.

c. Swiss franc and Japanese yen

The Japanese yen and Swiss franc are often treated as special currencies. In this section, we examine our model fit of these two currencies. While we still observe a rising explanatory power of the model over time, we note that some of the regressors have the wrong sign, so the model is not successful in explaining the exchange rates of dollar/Japanese yen and dollar/Swiss franc.

These two currencies are unique among the G10 for three reasons:

1. Japan and Switzerland both experienced prolonged stretches of negative inflation.
2. Both have undertaken large sterilized foreign exchange intervention.
3. Both are like the U.S., in that they are considered to be safe-haven currencies.

The baseline model does not account for these features. During periods of sustained deflation, the real interest rate and the measure of expected inflation (inflation over the previous 12 months) are unlikely to be good measures of the monetary stance. Sterilized intervention introduces another channel of influence on exchange rates that is not accounted for in the baseline model. And, our measure of global risk is introduced to capture the notion that the U.S. dollar is a safe haven during times of global financial stress, but the yen and Swiss franc are also considered safe havens during these times, so the variable may not be helpful in predicting changes in the bilateral dollar exchange rates with these two currencies.

In Table 5, we report the regression results of equation (1) for CHF and JPY exchange rates in the left panel and the variant of the equation with change of inflation in the right panel respectively. We observe that the R^2s are much lower for these regressions. They are about 0.10 for R^2 and 0.07 for adjusted R^2. The coefficients for the U.S. monetary variables – the real interest rate and inflation level have the predicted sign (negative) and are statistically significant. However, for other variables, the coefficient estimates are not always significant across different specifications and often feature the opposite sign as that predicted by theory. For example, the foreign interest rate effect is estimated to be negative for JPY on the left panel and CHF on the right panel, which is opposite the prediction of a monetary model. In addition, the $\Delta RISK_t$ variable is estimated to have a positive sign for the JPY case, which is opposite to what we observed in Table 1 across all other currencies. The coefficient of liquidity measure is still negatively significant for CHF but not for JPY. Finally, the lagged real exchange and TB/GDP variables are no longer statistically significant.

The particular ways in which the baseline model fails are indicative of the unique features of these three economies. The variables that measure the stance of U.S. monetary policy are successful, and the relative convenience yield correctly predicts the movement of the exchange rate. But the measures of the real interest rate and inflation expectations are not calibrated correctly to capture the monetary stance of these two countries that have had extended periods of negative inflation. The global risk variable is also statistically insignificant (or significant but with the "wrong" sign), which we would expect given the yen and Swiss franc, along with the U.S. dollar, are safe haven currencies. And, finally, the trade balance variable cannot be an accurate measure the predictions from the portfolio balance model, because sterilized intervention may significantly alter the supply of bonds of each currency that are outstanding.

d. Baseline model excluding the Global Financial Crisis

In Table 6, we look at the post-1999 regression exercise but exclude the sample period 2008-2009. We want to check to see whether the good fit of the model in the last quarter century is driven simply by the effect of the dollar fluctuation during the crisis.

Comparing with Table 1, the R^2s of the regressions reported in Table 6 are generally lowered by 3 to 11 percentage point. We observe that most of the results are robust to excluding the 2008-2009 period. Some of the foreign interest rate coefficients lose significance but the coefficient is still estimated to be very positively significant for the panel regressions.

In Figure 8, we produce the counterpart of Figure 1, but excluding the period 2008-2009 in the estimation process. That is, we do not estimate the model using 2008 to 2009 data and we do not fit the data in 2008-2009. The figure shows that the model fit across currencies is as good as Figure 1, except for GBP post 2009. This indicates that the good fit of the two figures is not coming from the dramatic 2008-2009 exchange rate movements in the crisis period. The poor fit of GBP in the post 2009 period is particularly driven by the 2016-2020 period when the GBP depreciated sharply. This is perhaps due to uncertainty arising from the Brexit Referendum and the follow-ups after the Brexit deal, which are unlikely to be well-captured by our simple model.

In Figure 9 and Figure 10, we conduct the 20-year rolling window regression without 2008-2009 and report the t-statistics and F-statistics, respectively. Again, we drop the liquidity variables because of lack of data availability. The end period becomes Aug 2001 in this case as we still maintain the sample window to be 240-months. Consistent with what we find in Figure 3 and Figure 4, we still find many right-hand-side variables become highly statistically significant in the second half of the period,[4] such as a positively significant change of foreign interest rate and foreign inflation level, negatively significant change of home interest rates, home inflation, U.S. trade balance to GDP and the change global risk measure $\Delta RISK_t$. The F-statistics and R^2 values are persistently rising since 1983. (Again, an exception is the U.K. pound, for which these statistics decreased in the 1990s before rising again at the end of the sample.) This shows that the significant results in the second half of the sample are not driven by the crisis.

In Figure 11, we re-examine the Meese Rogoff out-of-sample fit exercise without the 2008-2009 sample. We still observe that the Clark West statistics indicate the exchange rate model fits worse than the random walk in the early years and begins to outperform a random walk model in samples ending a in the latter half of the sample for most currency pairs. For the GBP case, there is some evidence that the exchange rate model beats a random walk model, but the Clark West statistics is not stably above the 1.65 dash line, which is again potentially influenced by the Brexit period.

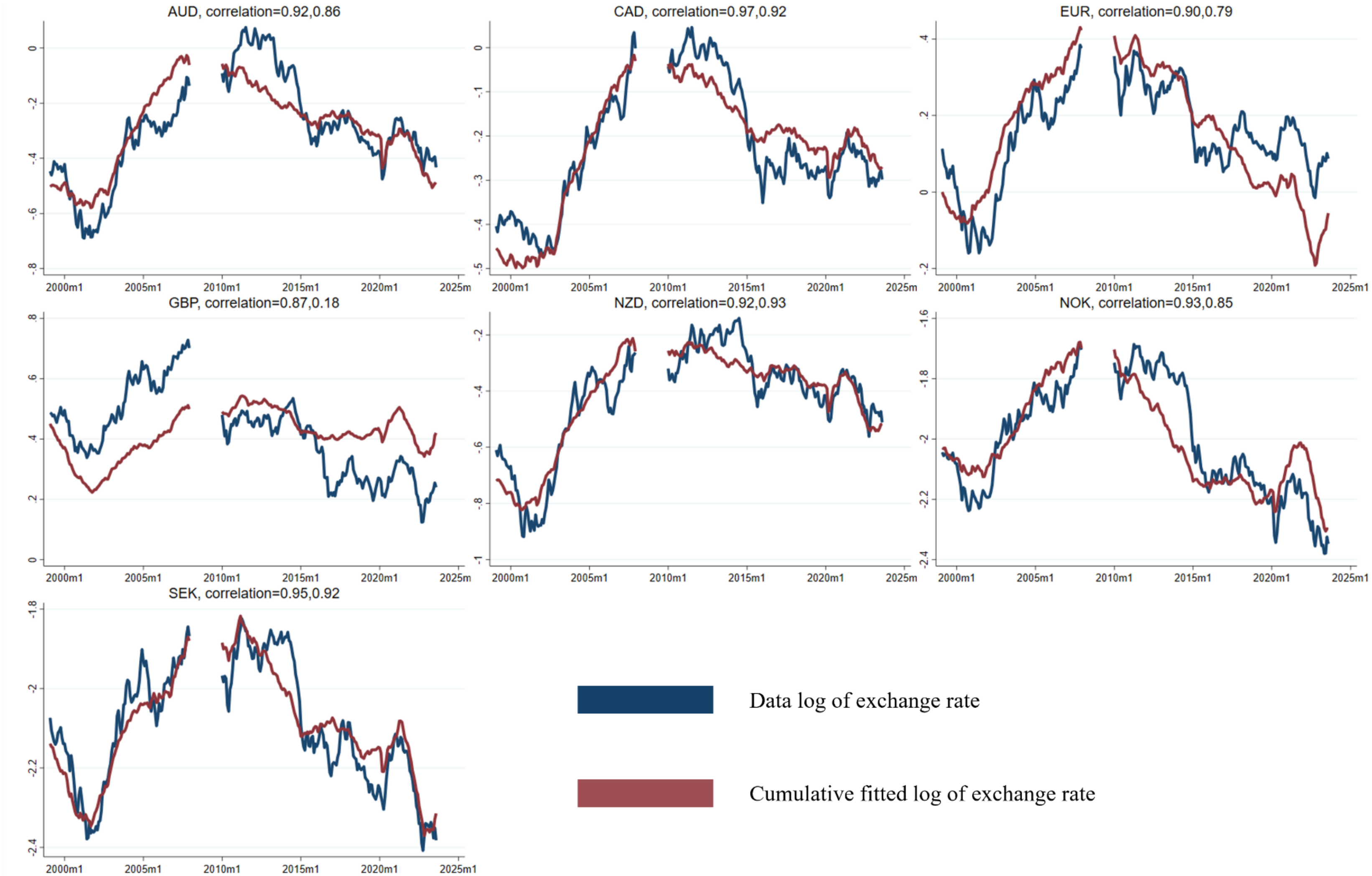
[4] The first half of the sample period (from 1973-1987) of these Figure 9 and Figure 10 are the same as Figure 3 and Figure 4 because the sample period of the rolling window regression do not include 2008-2009.

Table 6: Baseline regression with inflation level and convenience yield, excluding 2008-2009

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-1.04***	-1.76***	-2.39***	-1.16***	-0.79**	-1.83***	-1.86***	-1.37***	-1.38***
	(-3.25)	(-6.00)	(-7.27)	(-3.91)	(-2.16)	(-5.53)	(-5.70)	(-6.13)	(-6.19)
Δr_t^*	0.79***	1.17***	2.12***	1.11***	0.60	0.09	0.48	0.68***	0.69***
	(2.78)	(4.20)	(5.13)	(2.73)	(1.61)	(0.43)	(1.41)	(4.50)	(4.61)
π_t	-0.21*	-0.16	-0.57***	-0.50***	-0.49***	-0.14	-0.61***	-0.33***	-0.33***
	(-1.96)	(-1.21)	(-3.81)	(-3.80)	(-3.16)	(-1.14)	(-4.79)	(-4.04)	(-4.06)
π_t^*	0.03	0.12	0.47***	0.28**	0.29*	-0.27*	0.27***	0.16**	0.19***
	(0.20)	(0.72)	(3.45)	(2.43)	(1.91)	(-1.82)	(2.88)	(2.40)	(3.04)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01*	-0.01***	-0.03***	-0.03***	-0.02***	-0.02***	-0.02***
	(-8.07)	(-7.31)	(-1.80)	(-2.65)	(-5.76)	(-6.19)	(-4.02)	(-6.39)	(-6.15)
q_{t-1}	-0.01	-0.01	-0.02*	-0.03**	-0.01	-0.03***	-0.00	-0.01*	-0.00
	(-1.22)	(-1.24)	(-1.90)	(-2.53)	(-1.49)	(-2.85)	(-0.47)	(-1.72)	(-0.17)
$\frac{TB}{GDP_t}$	-0.46***	-0.46***	-0.59***	-0.83***	-0.38*	-0.58***	-0.75***	-0.54***	-0.50***
	(-2.70)	(-3.76)	(-3.63)	(-4.00)	(-1.93)	(-2.72)	(-3.84)	(-4.21)	(-3.93)
$\Delta\eta_t$	-1.49	-2.51***	-1.95*	-1.77*	-1.59**	-1.32	-2.02**	-1.58**	-1.64**
	(-1.54)	(-2.62)	(-1.90)	(-1.94)	(-1.98)	(-1.64)	(-2.53)	(-2.22)	(-2.33)
N	272	272	271	272	272	272	272	1903	1903
F	11.96	13.09	10.34	5.89	6.88	10.68	8.97	15.45	14.47
R2	0.27	0.28	0.24	0.15	0.17	0.25	0.21		0.17
R2_adj	0.24	0.26	0.22	0.13	0.15	0.22	0.19		
R2_within								0.17	

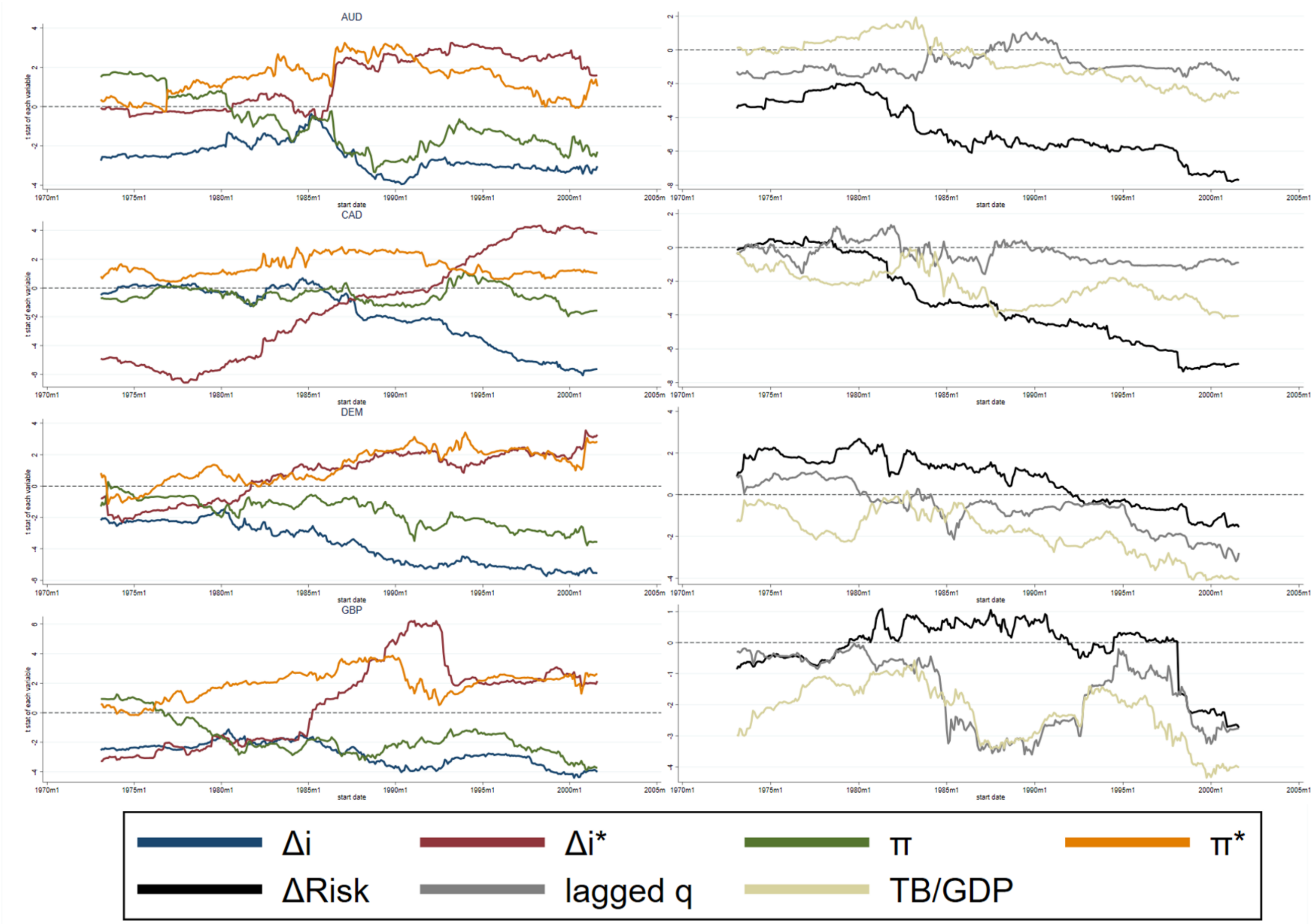
Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023 but excludes 2008-2009. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the measure of the U.S. convenience yield relative to the foreign country, using 1-year government bond and rates, as in Engel and Wu (2023).

Figure 8: Data and model implied exchange rates (excluding 2008-2009)



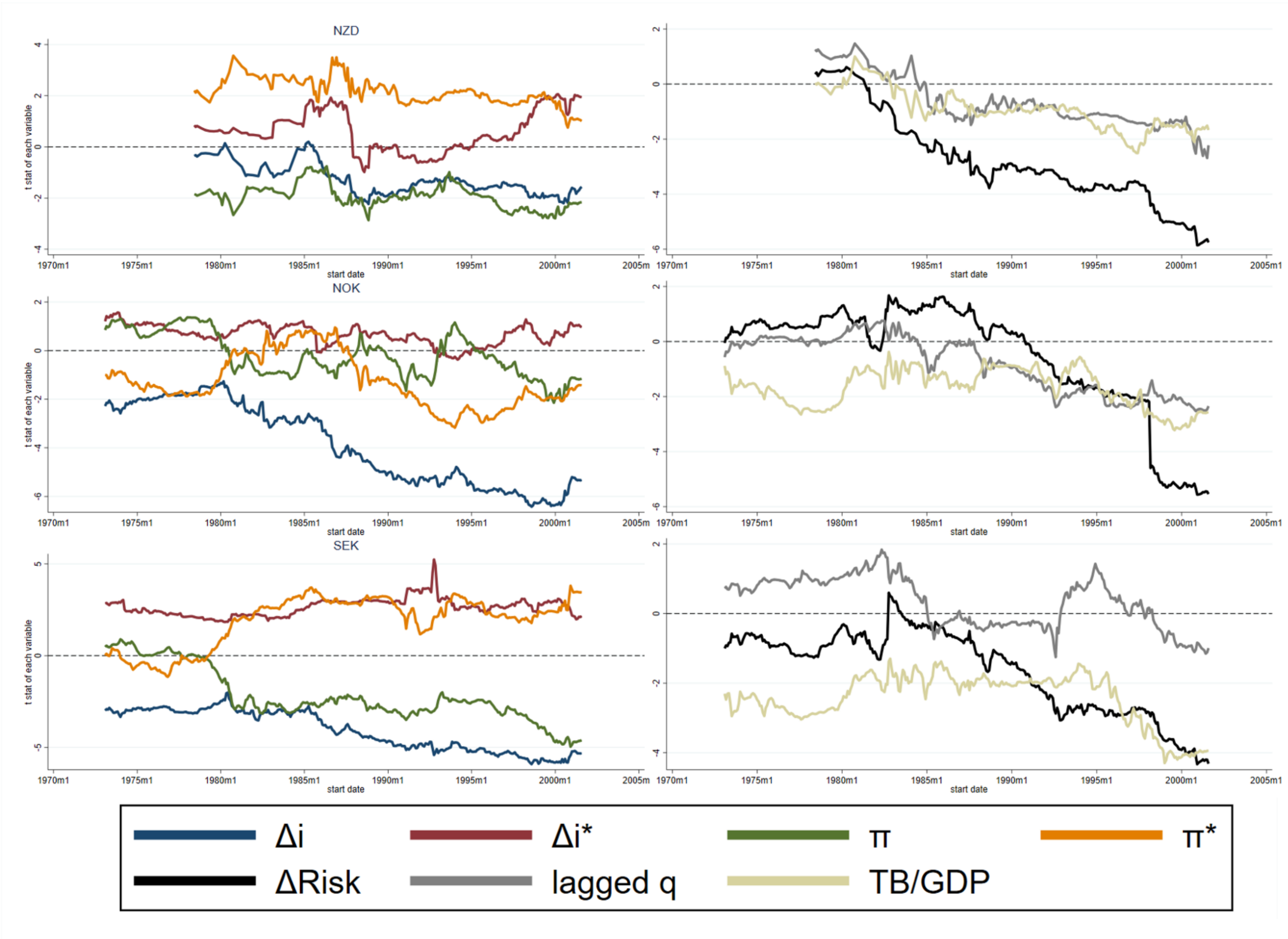
Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023 and excludes 2008-09. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the data series. Correlations of the two series are reported in the subtitles. Exchange rate is defined as the U.S. dollar price of a foreign currency, an increase in value is a U.S. dollar depreciation.

Figure 9: *t* statistics of 20-year rolling window regressions of equation (1), excluding 2008-2009



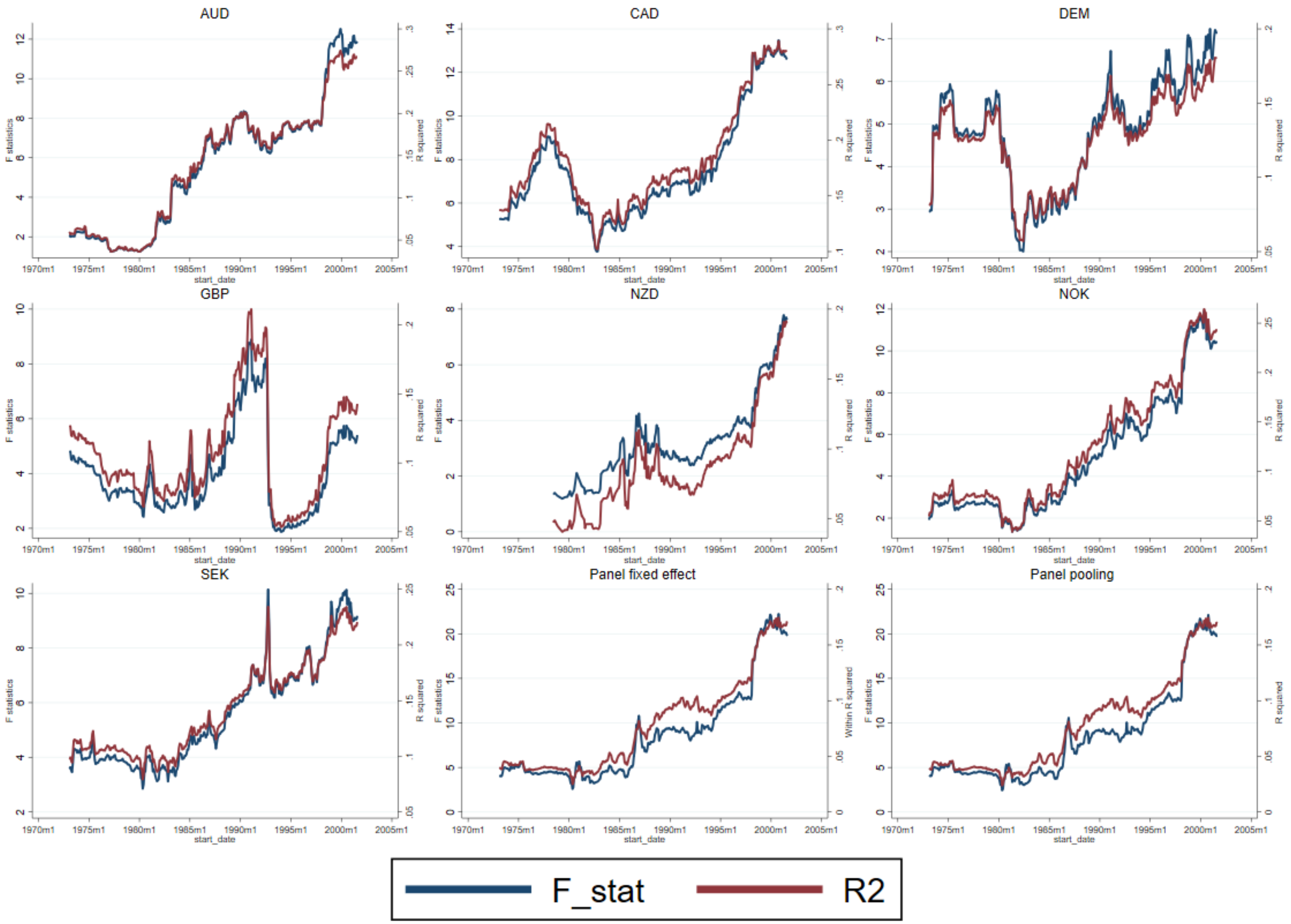
Note: The figure reports the t-statistics of each of the variables in equation (1) with a 20-year rolling window regression. X-axis correspond to the start date of the rolling window regression. Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. The last regression is Sep 2001-Aug 2023.

Figure 9: t statistics of 20-year rolling window regressions of equation (1), excluding 2008-2009, continued



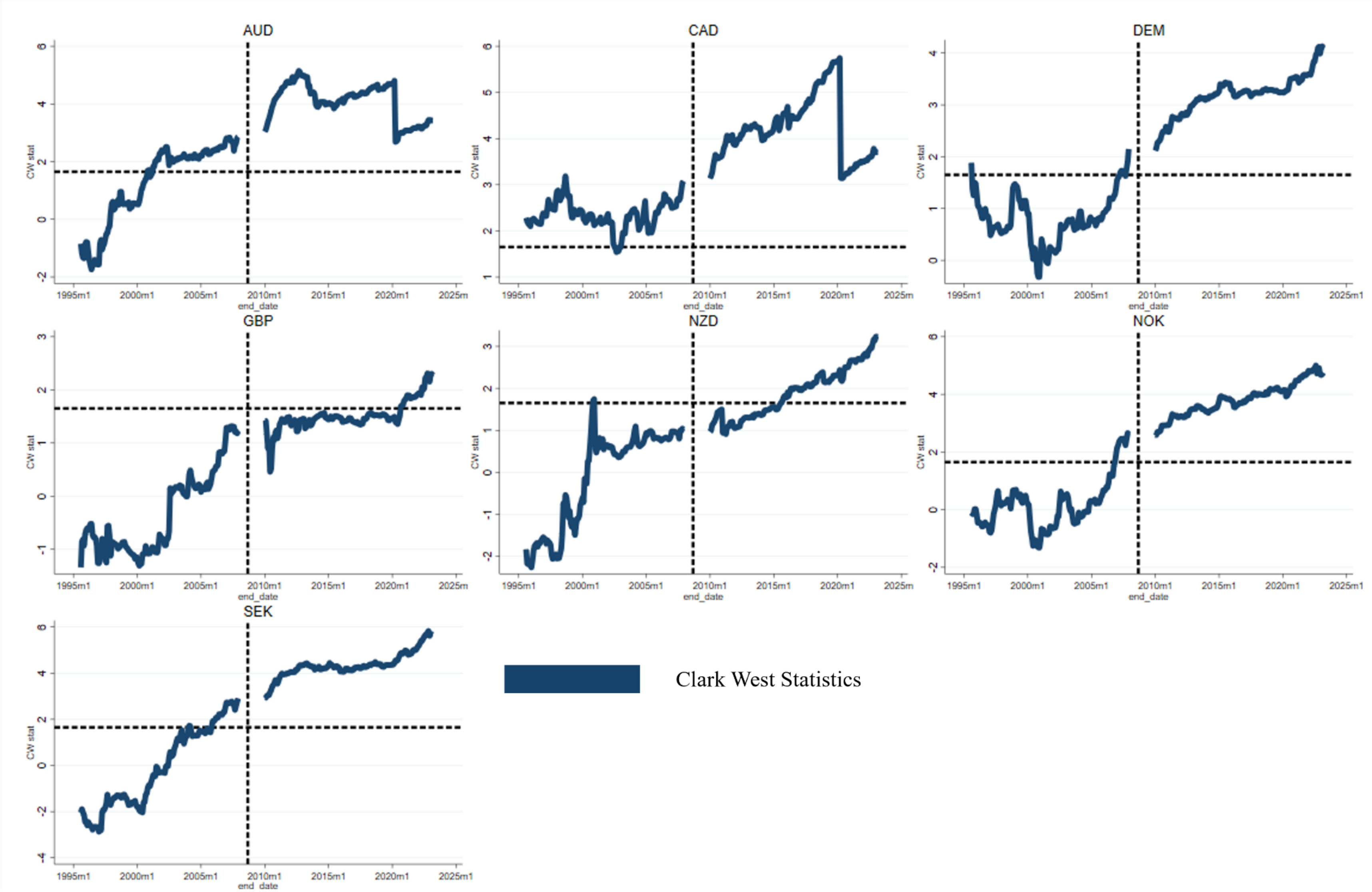
Note: The figure reports the t-statistics of each of the variables in equation (1) with a 20-year rolling window regression. X-axis correspond to the start date of the rolling window regression. Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2001-Aug 2023.

Figure 10: F-statistic and R^2 of 20-year rolling window regressions of equation (1), excluding 2008-2009



Note: The figure reports the F-statistics in equation (1) with a 20-year rolling window regression. X-axis corresponds to the start date of the rolling window regression. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2001-Aug 2023.

Figure 11: Clark West statistics of out-of-sample fit (excluding 2008-2009)



Note: The figure reports the Clark West statistics of each of the variables in equation (1) with 20-year rolling window regressions, excluding sample from 2008 to 2009. X-axis corresponds to sample end period of the Clark West statistic. The first Clark West statistic is computed with prediction errors from Mar 1993 to Sep 1995. The sample of prediction errors keep increasing till the end of sample. The last Clark West statistic is computed with prediction errors from Mar 1993 to Aug 2023.

4. Why the model didn't work in the old days

We find that a standard exchange-rate model fits U.S. dollar exchange rates well, especially when augmented by measures of global risk and liquidity demand. However, we see that the model did not fit well in the 1970s and 1980s, and only gradually began to perform better beginning in the 1990s. We contend that the change that occurred resulted from a regime-shift in monetary policy. In the early years of our sample, monetary policy in the U.S. and other countries was insufficiently responsive to inflation. The failure of the stability condition – the Taylor principle that real interest rates must ultimately rise sufficiently in response to inflation – may introduce a role for a self-fulfilling expectation indeterminacy into real and nominal exchange rates.

We first recap the theoretical result in the context of a simple open-economy model. With this in hand, we briefly draw some contrasts with the "sunspot" model and "noise trading" models. Then, in the next sub-section, we present some evidence that the closer relation of the exchange rate to its fundamental determinants coincides with increasing credibility of monetary policy.

a. Model with self-fulling expectations

We illustrate how equilibria with self-fulfilling expectations can emerge when the Taylor Principle is not satisfied in a simple two-country open-economy New Keynesian model. Such a model is necessarily simplified relative to the more detailed general equilibrium models in the literature that determine exchange rates but is useful for illustrative purposes.

We use the "canonical" three-equation model from Engel's (2014) *Handbook* chapter. We augment that model with shocks are added to the equation for relative returns (uncovered interest parity) and the relative Phillips curve. The solution presented here is the general one, rather than the special case in Engel (2014), so that we can clearly demonstrate how a self-fulfilling prophecies might be possible.

The model consists, first, of an equation that has the difference in the expected return on a foreign interest-earning asset and the home interest-earning asset (adjusted for exchange rate expectations) is equal to some exogenous random variable. This exogenous variable might be a risk premium or liquidity premium, for example, that would be endogenous in a more sophisticated set-up. The second equation is a version of the home minus foreign Phillips curve, which Engel (2014) notes can be derived under special symmetry assumptions about the home and foreign economies.[5] The third equation is given by

[5] Some of the specifics of the derivation and the dynamics of exchange rates under this system and closely related alternative set-ups are derived more completely in Engel (2019).

the home minus foreign monetary policy rules, in which the interest rate in each country is set to target inflation in that country, but with some interest-rate smoothing.

These three equations are, respectively,

(2) $$E_t s_{t+1} - s_t + i_t^R = \rho_t$$

(3) $$\pi_t^R = \delta\left(q_t - \tilde{q}_t\right) + \beta E_t \pi_{t+1}^R$$

(4) $$i_t^R = \sigma \pi_t^R + \alpha i_{t-1}^R + \varepsilon_t^R$$

A superscript R in these equations refers to the home country minus foreign country variable. s_t is the log of the exchange rate, home currency per unit of foreign currency. The relative nominal interest rate is i_t^R. ρ_t is the expected excess return on the foreign bond (perhaps, for example, a foreign exchange risk premium), and is exogenous. Equation (2) is the financial market equilibrium condition.

Equation (3) represents the evolution of home relative to foreign inflation. π_t^R is the relative inflation rate. It is related to the log of the real exchange rate, q_t, expressed as the relative price of foreign goods to home goods. The exogenous term $\tilde{q}_t$ represents factors that drive the "equilibrium" real exchange rate in the long run, such as total factor productivity or demographics. The equation says that home inflation tends to increase when the real exchange rate is above its long-run value, meaning that home prices rise to "catch up" with foreign prices. As in forward-looking prices of staggered price setting, expected future inflation also matters for current inflation rates.

The third equation represents the rule for setting monetary policy. In each country, policymakers target their own inflation rate, using the interest-rate instrument. But the interest rate adjustment occurs gradually, which is represented by the lagged interest rate term in the rule. The final term, ε_t^R, is exogenous and stands for other factors that might influence monetary policy.

We can gather these equations and write the system as:

(5) $E_t x_{t+1} = B x_t + \eta_t$, where

$$x_t = \begin{bmatrix} \pi_t^R \\ q_t \\ i_{t-1}^R \end{bmatrix}, \qquad \eta_t = \begin{bmatrix} \dfrac{\delta}{\beta}\tilde{q}_t \\ \rho_t - \dfrac{\delta}{\beta}\tilde{q}_t + \varepsilon_t^R \\ \varepsilon_t^R \end{bmatrix}, \qquad B = \begin{bmatrix} 1/\beta & -\delta/\beta & 0 \\ (\sigma\beta - 1)/\beta & (\beta+\delta)/\beta & \alpha \\ \sigma & 0 & \alpha \end{bmatrix}.$$

Decompose the B matrix as $B = A^{-1}\Lambda A$, where Λ is the diagonal matrix whose elements are the eigenvalues of B, and A is the matrix of row eigenvectors. Then we can pre-multiply (5) by the matrix A to write the system as:

(6) $$E_t z_{t+1} = \Lambda z_t + A\eta_t \text{, where } z = Ax.$$

This gives us a system of three univariate difference equations. One can easily show that if $\sigma + \alpha > 1$, two of the eigenvalues of B are greater than one, and one is less than one. The system has one predetermined variable at time t (i_{t-1}), and two variables (π_t^R and q_t) that can jump at time t in response to shocks to the exogenous variables. In this case, the condition that $\sigma + \alpha > 1$ is the stability condition, and we can solve one of the elements of z_t "backwards" and two "forward". That is, we find:

(7) $$z_{1,t+1} = \lambda_1 z_{1,t} + A_1\eta_t,$$

where λ_1 is the eigenvalue that is less than one, and A_1 is the first row of the matrix A. We also have:

(8) $$z_{j,t} = -E_t \sum_{i=0}^{\infty} \left(\frac{1}{\lambda_j}\right)^i A_j \eta_{t+i} \qquad j = 2,3,$$

where λ_2 and λ_3 are the eigenvalues that are greater than one. In deriving (8), we have imposed the "no-bubbles" condition, $\lim_{i\to\infty} E_t \left(\frac{1}{\lambda_j}\right)^i A_j \eta_{t+j} = 0$. The solution for the real exchange rate and the other two variables can then be recovered from (7) and (8) by premultiplying (6) by A^{-1}.

It is well known that a system such as this does not have a unique solution (even with the no-bubbles condition imposed) and admits the possibility of extraneous variables influencing all the variables in the system. Suppose the parameters are at the border of stability, so $\sigma + \alpha = 1$. Two of the roots are given by:

$$\lambda_1 = \frac{1+\sigma+\beta(1-\sigma) - \sqrt{\left(1+\sigma+\beta(1-\sigma)\right)^2 + 4\beta(1-\sigma\delta)}}{2\beta} > \frac{1+\delta}{\beta} > 1,$$

$$\lambda_2 = \frac{1+\sigma+\beta(1-\sigma)+\sqrt{\left(1+\sigma+\beta(1-\sigma)\right)^2+4\beta(1-\sigma\delta)}}{2\beta} > \frac{1+\delta}{\beta} > 1.$$

and the third is given by $\lambda_3 = 1$.

We could use equations (7) and (8) to solve for z_{1t} and z_{2t}, but in the case of $\sigma+\alpha=1$, we cannot do the infinite forward iteration for z_{3t}, corresponding to the root $\lambda_3 = 1$. Instead, consider iterating that equation forward for *k* periods. We find:

(9) $$z_{3,t} = -E_t \sum_{i=0}^{k} A_3 \eta_{t+i} + E_t z_{3,t+k+1}.$$

If $z_{3,t}^0$ is one candidate solution that satisfies (9), then so is $z_{3,t}^1 = z_{3,t}^0 + \chi_t$, where χ_t is any random variable that satisfies $E_t \chi_{t+k} = \chi_t$. If the extraneous sunspot variable χ_t affects z_{3t}, then it affects the solution to all variables in the system when (6) is multiplied by A^{-1}. Since $q_t - q_{t-1}$, the change in the log of the real exchange rate is solved from this system, as is π_t^R, we can conclude that the change in the log of the nominal exchange rate, $s_t - s_{t-1} = q_t - q_{t-1} + \pi_t^R$, is also influenced by the sunspot.

The sunspot variable in this example could be any random variable that markets decide should affect exchange rates. It does not have to be an economic fundamental, though it could be. That is, aside from the "true" influence of the fundamentals on the exchange rate, there could be an additional effect if the markets choose a fundamental variable as a sunspot.

How do self-fulfilling expectations work here, and why do they not influence exchange rates when the stability condition is met? Suppose markets arbitrarily decide that home inflation should be greater. If the stability condition were satisfied, this could not happen. Higher home relative to foreign inflation would drive up the home relative real interest rate, and rational expectations would rule out the possibility of a purely expectations driven inflation rate currently. When the stability condition is met, markets recognize that the sunspot cannot cause the inflation because current and future monetary policy are strict enough to offset the impact on the economy. But when the stability condition is not met, a higher belief about inflation may be expected to be permanent because the monetary policy reaction is not strong enough to tame inflation. In turn, higher future expected inflation leads to higher inflation today, and a weaker currency. The beliefs that drive these expectations may be tied to an extraneous random variable, or to some macroeconomic variable that is fundamental.

Comparison to noise trader models

The sunspot model of exchange rates has some commonalities with "noise-trader" models such as Jeanne and Rose (2002), Devereux and Engel (2002), and Itskhoki and Mukhin (2021). In both cases, variables that are not economic fundamentals can play an important role in determining exchange rates. The relationship of the exchange rate to economic fundamentals may be obscured by the noise or sunspots, and in fact, the exchange rate could exhibit unpredictable, random-walk-like behavior.

There are some differences in the models. The model presented above does not require deviations from uncovered interest rate parity, while the aforementioned noise-trader models all rely on such a gap. However, this difference is not empirically relevant since it seems apparent that a risk premium and/or liquidity premium is one significant determinant of exchange rates, particularly in the past two decades.

Another difference is a modeling one: the sunspot model does not require the introduction of any private agents into a model that are not optimizing or do not have rational expectations. The failure of the stability condition lies at the feet of the monetary policymakers. In contrast, the noise-trader models require some agents that are financially large enough to play a role in determining exchange rates, yet who are not determining their currency trades based on an explicit rational optimizing criterion. There is the additional complexity of reconciling the existence of agents with such a large financial impact with their potential effect on other economic variables such as consumption and investment.

The most relevant difference is highlighted by the findings of the empirical section. The exchange rate models fit much better than they used to. We explain that change by a verifiable change in regime for monetary policy. It is more difficult to explain this with noise traders. The good fit of the models in the 21st century does not preclude the presence of noise trade. The fit of the models is not perfect (though we believe that it is not surprising the R^2 is less than one, because econometricians cannot exactly measure the forces in our model, such as expectations, risk and demand for liquidity.) Even if the fit is not perfect now, it is better than in the past. The noise trader approach requires some explanation for why the noise traders are less important than they used to be.

Relatedly, the sunspot model is consistent with Bacchetta and van Wincoop's (2004) scapegoat model. They find that different economic variables seem to be important for the exchange rate during different time periods pre-2000. In our telling of the story, the scapegoats are the different sunspots that the markets decide are important for determining exchange rates. This aspect of the data appears more difficult to reconcile with noise trading.

b. Evidence to support an inflation credibility channel

In this subsection, we provide three pieces of empirical evidence for the inflation credibility channel.

t-Statistics

The literature has produced evidence that the monetary policy rule for the U.S. began to shift during the Volcker era, so that Taylor rules estimated on data beginning in the mid-1980s offer support for monetary stability.[6] The advanced countries in our sample adopted inflation targeting a few years later: New Zealand in 1990, Canada in 1991, the U.K. in 1992, Sweden and Australia in 1993, Norway in 2001. One of the pillars of European Central Bank policy, beginning in 1999, is inflation targeting. Germany formally adopted inflation targeting in 1992 before the advent of the euro, though targeting inflation was always at the core of Bundesbank policy. The other two countries we have examined, Switzerland and Japan, are also somewhat exceptional in this regard as well. Both have had consistently low inflation, but Japan only formally adopted inflation targeting as an objective in 2013, and the Swiss National Bank never has.

The first evidence that lack of monetary policy credibility accounts for the poor fit of the model in the earlier part of the sample comes from the plots of the *t*-statistics for the monetary variables in the left-hand panel of Figure 2. The *t*-statistics are a useful measure of the contribution of the monetary variables precisely because they measure the orthogonal quantitative contributions of the variables to explaining exchange rate changes, weighted by the inverse of the precision of the parameter estimates.

First, consider the U.S. and foreign real interest rate. The U.S. real interest rate coefficient is generally estimated as negative throughout the exercises, but the first estimate uses data from March 1973 to February 1993, so already contains a significant bit of data in the post-Volcker era. Against the German mark, U.K. pound, Swedish krona and Norwegian krone, the coefficient is usually estimated to be negative and significantly so at the 5% level. A higher real U.S. interest rate also is generally associated with a stronger dollar against the Australian dollar, New Zealand dollar, and Canadian dollar, though the estimate is not always statistically significant. The overall impression is that during this period, current (as opposed to anticipated) U.S. monetary policy has the effect on exchange rates posited by the model, and that in part reflects the fact that inflation was much more directly targeted after Paul Volcker assumed the chairmanship of the Fed.

As we have noted above, the other countries only adopted inflation targeting beginning in the 1990s. With the exception of the Swedish krona, the coefficients on the real interest rates for those countries

[6] For example, see Clarida et al. (2000), Coibion and Goronichenko (2011), or Hirose et al. (2023).

are never estimated to be significantly positive in the earlier part of the sample. The model predicts a positive coefficient, so an increase in the real interest rate in those countries should lead, *ceteris paribus*, so a depreciation of the dollar. However, as we have noted, in the latter half of the sample, these coefficients are estimated to be positive, and significant.

The other monetary variable is our proxy for expected inflation. As the title of Clarida and Waldman's (2008) paper asks, "Is bad news about inflation good news for the exchange rate?" That paper makes the point that when the Taylor principle is satisfied, an increase in expected inflation in a country should appreciate the currency, in anticipation of future tighter monetary conditions. For the U.S., when policy is credible, the estimated coefficient should be negative. It is apparent that in the earlier part of the sample, inflation expectations did not definitively have this effect on the exchange rate. The parameter estimates are almost never significantly negative, and for all the currencies except the NZD, there are periods early in the data in which the estimate is positive. Yet, by the last window of estimation, the coefficients are all significantly negative. Even in the case of the U.S., high inflation over the previous 12 months did not convince markets that the Fed would be tightening, leading to a stronger dollar, until the latter half of the period.

The findings for the measure of expected inflation in the foreign countries is similar, except that it is not until later windows (compared to the coefficient on U.S. inflation) that the estimates become significantly positive as the model would predict (and, in the case of NOK, do not become significantly positive.) This again is consistent with the observation that these countries did not adopt an inflation-targeting monetary policy until somewhat later than the Fed in effect did.

By way of comparison, the noise-trader model (discussed above) would not predict a change in the goodness-of-fit of the monetary variables over this period.

A related piece of evidence about the importance of the monetary policy variables is presented in Appendix C Figure 1A, B and C. There we have re-estimated the model over the January 1999 – August 2023 sample but included only some of the explanatory variables in each of three regressions. We then construct fitted values as in Figure 1. In the first group, we use only the U.S. and foreign real interest rate, and the U.S. and foreign inflation variables to explain exchange rate changes; in the second group we use the risk, liquidity and portfolio variables (the spread, the convenience yield, and the trade balance); and the third uses only the lagged real exchange rate.

It is clear from these figures that the monetary policy variables are crucial in providing a good fit over this period. They appear to track most of the important movements in the value of the dollar. While the financial variables are certainly important, on their own, they do not capture all the major exchange rate movements. The large appreciation of the dollar at the time of the global financial crisis depends on these risk-related variables, but both the monetary and risk variables are needed to reproduce the close overall fit in Figure 1. The lagged real exchange rate is not helpful at all, which indicates that the overall good fit of the model from Figure 1 is not somehow "baked in" by including the lagged real exchange rate in the regression.

Estimated Taylor Rules with Switching

We present here direct evidence of the evolving monetary policy regimes in the U.S. and other countries. As we noted in the introduction, there is an extensive literature that investigated the increasing credibility of U.S. monetary policy in the late 20th century. Here, we develop a framework that we can apply consistently across all countries.

For each country, we examine a version of the Taylor rule for monetary policy estimated in Coibion and Gorodnichenko (2011) using quarterly data:

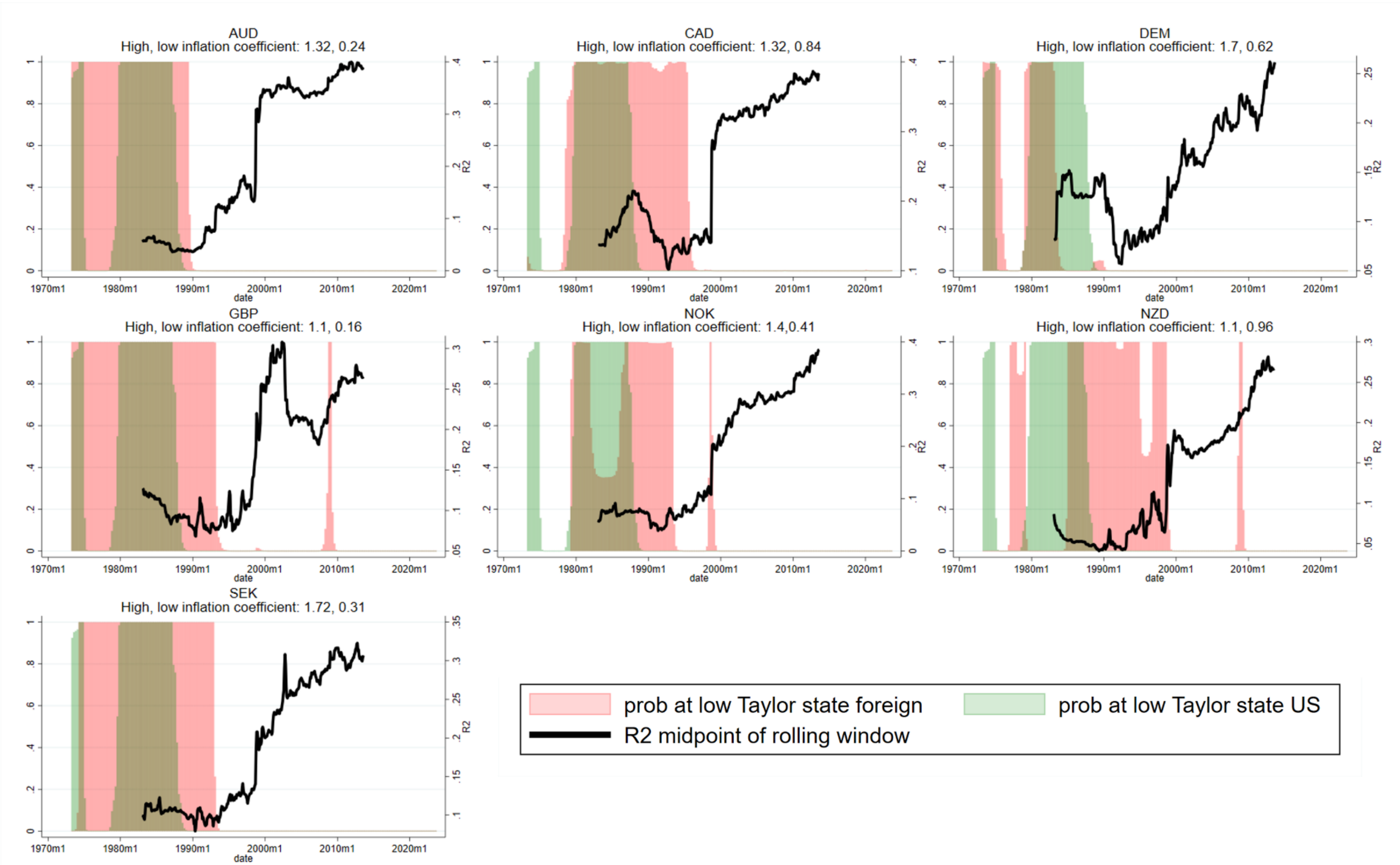
$$i_{j,t} = c_j + \left(1-\rho_j\right)\left(\phi_{\pi j}\pi_{j,t} + \phi_{gj}g_{j,t} + \phi_{xj}x_{j.t}\right) + \rho_j i_{j,t-1} + u_{j,t} \qquad (10)$$

For country j, $i_{j,t}$ is the policy rate at the end of the quarter (expressed as an annualized rate); $\pi_{j,t}$ is consumer price inflation over the past year; $g_{j,t}$ is the quarterly growth rate of real GDP; and, $x_{j,t}$ is a measure of the output gap. The output gap is estimated with an HP filter with parameter 1600 and re-estimated each period to allow for real time reassessment of the full-employment output level.

We are interested in the coefficient $\phi_{\pi j}$ - credible policy implies the stability condition will be satisfied. We estimate a switching model for the monetary rule (10) for each country in which the stable regime is an absorbing state. Specifically, we estimate a two-state Markov-switching model as in Hamilton (1989) but impose constraints. We constrain $\phi_{\pi j}$ to be greater than one in one state, and we constrain the probability of switching from the high state to the low state to be zero.[7] The model is estimated by maximum likelihood. The smoothed probabilities of the states in which $\phi_{\pi j} < 1$ are reported in Figure 12.

[7] Appendix D reports results when the stable state is not constrained to be an absorbing state. The general findings are not changed – that it is only in recent decades that monetary policy rules have conclusively satisfied the Taylor condition.

Figure 12: Probability of stability condition is satisfied from the Markov switching model



Note: The figure reports the probability of $\phi_{\pi j}$<1 (the condition that Taylor principle fails) for the US in the shaded green region and the foreign country of the title of each subfigure in the shaded red region. The overlapped area is shaded with brown color. The coefficients of $\phi_{\pi j}$ at the high and low Markov states are reported in the subfigure title. The black lines report the R^2 measure of 20 year rolling window regressions in equation (1). For each date, the black line indicates the midpoint of the rolling window.

In Figure 12, the shaded green areas, which are the same in all panels, are the probability that the state is such that $\phi_{\pi j} < 1$ when j = U.S. The shaded red areas give the probability that the non-U.S. country j is in the regime in which $\phi_{\pi j} < 1$. When both are in the regime of $\phi_{\pi j} < 1$, the probability is given by brown shading. Figure 12 also reports the estimated values of $\phi_{\pi j}$ in each regime.

Also plotted in Figure 12 is the R^2 value of the rolling regressions of the exchange-rate model. The plotted value corresponds to the mid-point of each 20-year sample used to estimate the exchange-rate model.

The graphs show that in each country, the Taylor-rule coefficient was less than one in the earlier part of the sample, then switched to being greater than one. The switch occurs later for the non-U.S. countries than for the U.S., which is consistent with the narrative that the U.S. monetary policy changed to a stable regime in the 1980s, while the other countries moved in that direction somewhat later. An exception is Germany, which, according to the estimates, appears to have switched to a credible regime in the early- to mid-1980s.

The fit of the exchange-rate model begins to increase with a lag of a few years after both countries in each pair have switched to the credible monetary policy regime. This gap might represent time for markets to learn and become convinced that the Taylor rules have changed. The evidence presented in the next section supports the notion that credibility continued to climb for the non-U.S. countries even in the early 21st century.

Response of Exchange Rate to Inflation Announcement

One clear indication of whether markets believe that central banks will react to higher inflation by raising real interest rates comes from the movements of the exchange rate in response to news about inflation. We make use of the event when national statistical agencies announce the measure of inflation in the country for an earlier month and investigate the exchange rate response on those dates. For example, in the middle of each month, the Bureau of Labor Statistics releases the measure of consumer price inflation in the U.S. for the previous month. The “news” or “surprise” about inflation can be measured by taking the difference between the actual inflation announcement and the predicted inflation number by surveys of financial market participants.

When central banks are following a credible inflation targeting policy, they are expected ultimately to increase real interest rates when inflation turns out to be higher than expected. In that case, the country's currency should appreciate at the time of the announcement.[8]

We obtained Bloomberg's survey of cross-country consumer price inflation forecast by financial market participants a few days prior to the announcement, asking their expectation of what the announced inflation will be. Note this survey does not ask for a forecast of inflation – it asks for a prediction about what the statistical authorities will reveal about past inflation.

For our purposes, however, the survey data is not ideal because Bloomberg only began to conduct the surveys for our set of countries in 1997. So, we cannot see the sharp differences that might prevail in the years previous to inflation targeting compared to the more recent years. However, we can at least see if the credible policy implied relationship between inflation surprise and exchange rate holds post-1997.

In Table 7, we report the coefficient estimates of regressing daily exchange rate change on the inflation surprise measures of the U.S. and foreign countries on the inflation announcement dates. Formally, the regression is:

$$\Delta s_t = \alpha + \beta(inflation\ annoucnement\ value - inflation\ survey\ value)_t + u_t$$

In panel A, the sample period is January 1997 to April 2024.[9] As discussed above, we exclude the sample from 2008-2009 to avoid the results from being driven by the Global Financial Crisis. The coefficients estimated confirm our hypothesis. When the foreign country inflation announced is higher than the market expectation, there is a U.S. dollar depreciation relative to the foreign currency, indicating the market is expecting the foreign government to tighten the future real interest rates. A coefficient of one in this regression indicates the U.S. dollar depreciates by one percent when the annualized inflation surprise is one percentage point. On the other hand, when the U.S. inflation surprise is positive, the U.S. dollar tend to appreciate during the inflation announcement dates.

While we do not directly observe the inflation surprise measure pre-1997 due to data availability, we can compare our results with the literature findings. Following the spirit of Anderson, et al. (2003), multiple papers document the exchange rate response to foreign macro news announcements and especially inflation news announcements. These papers include Cheung et al (2019), Hutchison and Sushko (2013), Kim (1998), Mazigi (2002), Clare and Courtenay (2001) and Joo et al (2009). However,

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[8] See Clarida and Waldman (2008). Closely related predecessors to this study are Engel and Frankel (1984) and Anderson, et al. (2003).

[9] The exact start date differs by currency. Appendix Table reports sample period by currency. The exchange rate we use here is the New York noon time exchange rate. We adjusted for the time zone difference if the inflation announcement is made after New York noon time.

in contrast to what we find in Panel A of Table 7, these papers do not find a significant relationship between foreign inflation surprise and exchange rate adjustment around the short window in the 1990s. This indicates a change of the statistical relationship between the two and it coincides with the gradual increase in inflation credibility as suggested in Figure 12.

In panel B, we focus on the sub-sample of 2010 to April 2024. The coefficient estimates are generally improved in this sub-period. For the panel regressions, the coefficients are at least one standard deviation higher than the original estimates. For individual currency, there are visible improvement for the foreign inflation surprise for Canadian dollar and Norwegian Krone. Especially for Euro, the foreign surprise coefficient turns to positive. The US inflation surprise coefficient estimates also improved. Overall, the relationship of inflation surprise and exchange rate holds more tightly and indicates a slow gain process in the second half of the sample.

To summarize, the empirical evidence provided in this subsection suggests that the Taylor principle is not satisfied for a large set of country pre-2000s. This reflects as a weak relationship between exchange rate and fundamental in the early sample. But the improvement in inflation credibility from post-2000s sample is associated with a strong connection of exchange rate with monetary variables and especially inflation surprise in a high frequency setting.

5. Conclusions

We find that the "standard" monetary model of exchange rates, augmented with the measures of global financial stress emphasized in recent studies does a good job in accounting for U.S. dollar exchange rate movements. The regression fits well, even when it is estimated on monthly changes in the log of exchange rates. The fit is good both in-sample, and out-of-sample. This is a change from the "old days" when the model fit poorly, which we confirm in our rolling regressions. We attribute the improvement in fit to the increasing credibility of monetary policy in the U.S. and other countries.

Admittedly, the fit is not perfect, in that the R^2 values, while quite high, are not 1.0. In part, that is surely due to the impossibility of perfectly measuring any of the variables in the model – inflation expectations, expectations of future monetary policy, the effect of supplies of currency-denominated debt, global risk and liquidity, as well as the equilibrium real exchange rate. There are probably other variables that have yet to be discovered by theory, or that have already been discovered but not included here, or that will become important in the future. Still, the models fit better than you might have thought.

Table 7: Daily exchange rate regression on the inflation announcement dates

	Panel fixed	AUD	CAD	EUR	GBP	NZD	NOK	SEK
Panel A: Full sample								
Foreign	0.76***	1.44***	0.64***	-0.93	0.68***	1.11***	0.63***	0.87***
surprise	(0.08)	(0.33)	(0.13)	(0.68)	(0.20)	(0.37)	(0.16)	(0.24)
U.S.	-0.61**	-0.70**	-0.38*	-0.56**	-0.48*	-0.78**	-0.60*	-0.79**
surprise	(0.28)	(0.35)	(0.21)	(0.26)	(0.25)	(0.36)	(0.34)	(0.32)
R^2	0.03	0.06	0.05	0.01	0.03	0.03	0.03	0.03
N	3490	404	581	469	560	381	552	543
Panel B: Post 2009 sample								
Foreign	0.85***	1.47***	0.83***	0.30	0.61**	1.07***	0.70***	1.14***
surprise	(0.10)	(0.45)	(0.18)	(1.02)	(0.25)	(0.40)	(0.19)	(0.31)
U.S.	-0.99**	-1.00**	-0.41	-1.21***	-0.95***	-0.95**	-1.03**	-1.38***
surprise	(0.39)	(0.44)	(0.28)	(0.32)	(0.35)	(0.45)	(0.47)	(0.44)
R^2	0.03	0.06	0.05	0.01	0.03	0.03	0.03	0.03
N	2081	229	338	315	325	223	327	324

Standard errors in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to April 2024 but excludes 2008-2009. The explanatory variable in all regression is the daily change of U.S. exchange rate with the currency in the column head on the inflation announcement dates. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. Panel fixed effect regressions and panel pooling regressions produce exactly the same coefficients since all the country fixed effects are almost zero. For space reason, we do not report the panel pooling regressions.

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Appendix A: Data

Unless otherwise specified, the sample period starts from March 1973 and ends in August 2023

Variables	Data source	Note
Exchange rates	Federal Reserve Economic Database (FRED)	FRED: DEXUSAL (AUD), DEXCAUS (CAD), EXGEUS (DEM), DEXUSEU (EUR), DEXJPUS (JPY), DEXNOUS (NOK), DEXUSNZ (NZD), DEXSDUS (SEK), DEXSZUS (CHF), DEXUSUK (GBP). We scale the DEM exchange rate such that the EUR and DEM exchange rate are the same in Jan 1999. Monthly average series are used.
Inflations	International Monetary Fund International Financial Statistics (IMF IFS)	Raw consumer price index data are downloaded. Inflation is computed as: $\log(CPI_t) - \log(CPI_{t-12})$. For Australia and New Zealand, because the CPI are reported quarterly, we use the last available CPI data for the whole quarter.
Nominal interest rates	Global Financial Database (GFD) and FRED	GFD: ITAUS3D (AUD), ITCAN3D (CAD), ITDEU3D (DEM), ITJPN3D (JPY), ITNOR3D (NOK), ITNZL3D (NZD), ITSWE3D (SEK), ITCHE3D (CHF), ITGBR3D (GBP). New Zealand's interest rate is only available from 1978. FRED: DTB3 (US)
Risk variables	FRED. For Gilchrist and Zakrajsek (2012) spreads, update and maintained by https://www.federalreserve.gov/econres/notes/feds-notes/updating-the-recession-risk-and-the-excess-bond-premium-20161006.html	FRED: AAAFF, BAAFF, AAA10Y, BAA10Y
Trade balance to GDP	FRED	BOPGSTB (post-1992), BOPBGS (pre-1992), GDP. Quarterly variables are interpolated.
Liquidity Ratio (Only used for post 1999 regressions)	FRED	sum of U.S. dollar financial commercial paper (FRED series: DTBSPCKFM) and short-term funding to U.S. banks is demand deposits (FRED series: DEMDEPSL) divided by the sum of reserves held at Federal Reserve banks and government securities held by commercial banks (the sum of TOTRESNS and USGSEC from FRED.)
Convenience yield/Liquidity yield	Datastream	The variable is constructed by using $f_t - s_t - (i_t - i_t^*)$ where f_t is one year forward and i_t, i_t^* are one year government bond interest rate. Datastream mnemonic for the Forward rates are:

(Only used for post 1999 regressions)		USAUDYF, USCADYF, USDEMYF, USJPYYF, USNZDYF, USSEKYF, USCHFYF, USNOKYF, USGBPYF. Datastream mnemonic for the 1 year government rates are: TRAU1YT, TRCN1YT, TRBD1YT, TRJP1YT, TRNZ1YT, TRNW1YT, TRSD1YT, TRSW1YT, TRUK1YT, TRUS1YT
Inflation expectation measures	Data shared by Benigno and Eggertson (2023)	Sample period: 1960Q1-2023Q2. Interpolated to quarterly frequency.

Appendix B: Model Estimated 1999:1-2023:8 without Liquidity Variables

Table B1 with inflation change and 3m Govt rate

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-3.92***	-3.85***	-3.34***	-2.41***	-4.12***	-2.53***	-3.19***	-2.74***	-2.72***
	(0.742)	(0.667)	(0.688)	(0.675)	(0.870)	(0.757)	(0.759)	(0.627)	(0.626)
Δr_t^*	4.37***	3.42***	2.73***	3.45***	4.10***	0.00	2.57***	1.31**	1.36**
	(0.794)	(0.714)	(0.821)	(0.763)	(0.860)	(0.288)	(0.869)	(0.631)	(0.644)
$\Delta \pi_t$	-3.41***	-2.62***	-1.59**	-1.55**	-4.06***	-0.65	-1.87**	-1.65**	-1.67**
	(0.812)	(0.725)	(0.775)	(0.752)	(0.955)	(0.812)	(0.836)	(0.694)	(0.691)
$\Delta \pi_t^*$	3.52***	2.65***	0.44	1.97**	3.54***	-0.50	1.80**	0.54	0.61
	(0.819)	(0.730)	(0.953)	(0.888)	(0.903)	(0.377)	(0.890)	(0.662)	(0.674)
$\Delta RISK_t$	-0.04***	-0.03***	-0.01***	-0.02***	-0.03***	-0.03***	-0.02***	-0.03***	-0.03***
	(0.004)	(0.002)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)
q_{t-1}	-0.00	-0.01*	-0.02**	-0.02	-0.01	-0.01	-0.01	-0.01*	0.00
	(0.006)	(0.007)	(0.009)	(0.011)	(0.008)	(0.008)	(0.009)	(0.006)	(0.000)
$\frac{TB}{GDP_t}$	-0.28*	-0.33***	-0.37**	-0.31*	-0.14	-0.43**	-0.33*	-0.31***	-0.27**
	(0.156)	(0.108)	(0.151)	(0.173)	(0.179)	(0.177)	(0.178)	(0.116)	(0.115)
N	296	296	295	296	296	296	296	2071	2071
F	28.13	25.02	11.57	12.03	14.00	16.23	12.00	24.09	23.27
R2	0.41	0.38	0.22	0.23	0.25	0.28	0.23		0.23
R2_adj	0.39	0.36	0.20	0.21	0.24	0.27	0.21		
R2_within								0.23	

Note: Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. $\mathrm{r_t}$ and $\mathrm{r_t^*}$ are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S

Table B2 with inflation level and 3m Govt rate

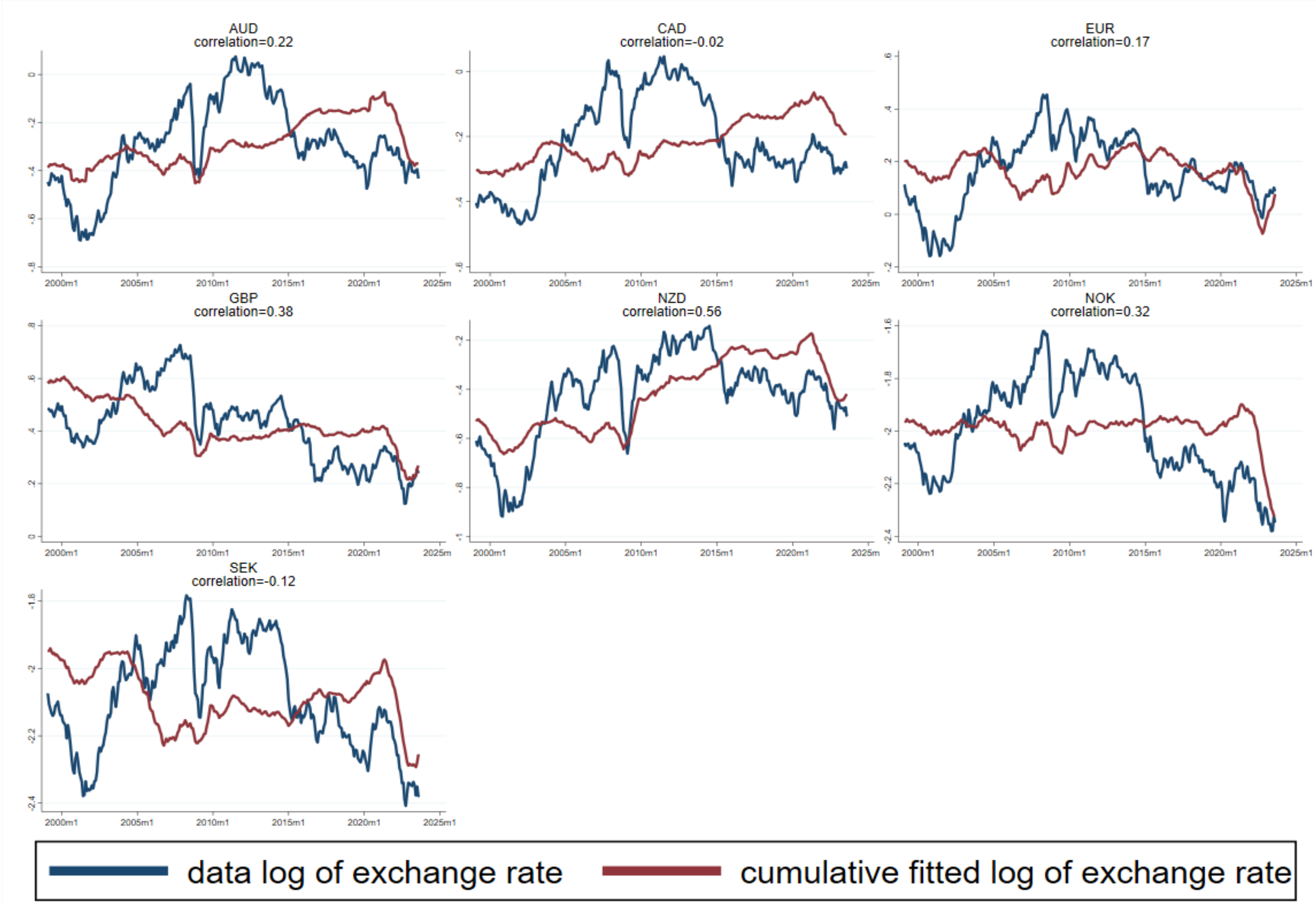
	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-1.11***	-1.64***	-2.33***	-1.37***	-0.93***	-1.88***	-1.88***	-1.47***	-1.46***
	(0.293)	(0.264)	(0.289)	(0.259)	(0.327)	(0.291)	(0.295)	(0.200)	(0.200)
Δr_t^*	1.06***	1.15***	2.21***	1.79***	1.02***	0.21	0.77**	0.89***	0.91***
	(0.279)	(0.272)	(0.394)	(0.373)	(0.358)	(0.199)	(0.333)	(0.168)	(0.169)
π_t	-0.26**	-0.22*	-0.69***	-0.34***	-0.46***	-0.22**	-0.57***	-0.35***	-0.34***
	(0.104)	(0.127)	(0.141)	(0.116)	(0.129)	(0.111)	(0.125)	(0.080)	(0.079)
π_t^*	0.04	0.15	0.53***	0.15	0.25*	-0.19	0.26***	0.16**	0.19***
	(0.127)	(0.156)	(0.131)	(0.108)	(0.137)	(0.130)	(0.095)	(0.069)	(0.065)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01***	-0.01***	-0.03***	-0.03***	-0.02***	-0.02***	-0.02***
	(0.003)	(0.002)	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
q_{t-1}	-0.01	-0.01	-0.02***	-0.04***	-0.01	-0.03***	-0.01	-0.01**	-0.00
	(0.007)	(0.007)	(0.009)	(0.012)	(0.008)	(0.009)	(0.010)	(0.006)	(0.000)
$\frac{TB}{GDP_t}$	-0.47***	-0.46***	-0.63***	-0.75***	-0.37*	-0.66***	-0.80***	-0.54***	-0.47***
	(0.173)	(0.124)	(0.163)	(0.211)	(0.196)	(0.210)	(0.201)	(0.126)	(0.124)
N	296	296	295	296	296	296	296	2071	2071
F	24.01	22.70	15.09	12.67	12.17	18.61	14.58	23.50	22.25
R2	0.37	0.36	0.27	0.24	0.23	0.31	0.26		0.24
R2_adj	0.35	0.34	0.25	0.22	0.21	0.29	0.24		
R2_within								0.24	

Note: Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S.

Appendix C: Figure 1 with only a subset of variables

Figure 1A: Fitted value generated from regression only including change of real interest rates and inflation level

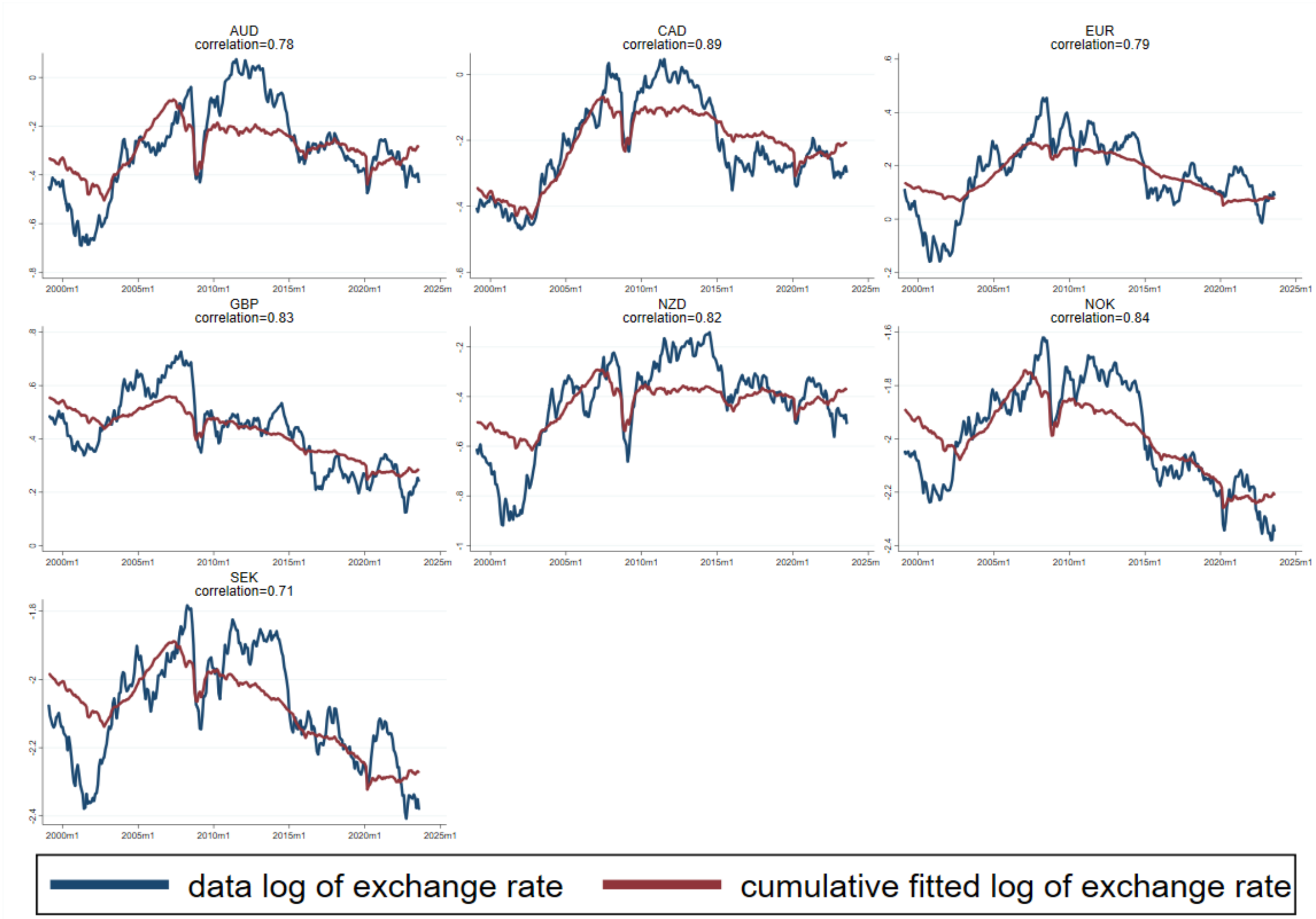
$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + u_t$$



Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the data series. Correlations of the two series are reported in the subtitles. The exchange rate is defined as the U.S. dollar price of a foreign currency: an increase in value is a U.S. dollar depreciation.

Figure 1B: Fitted value generated from only risk variable, TB and convenience yield

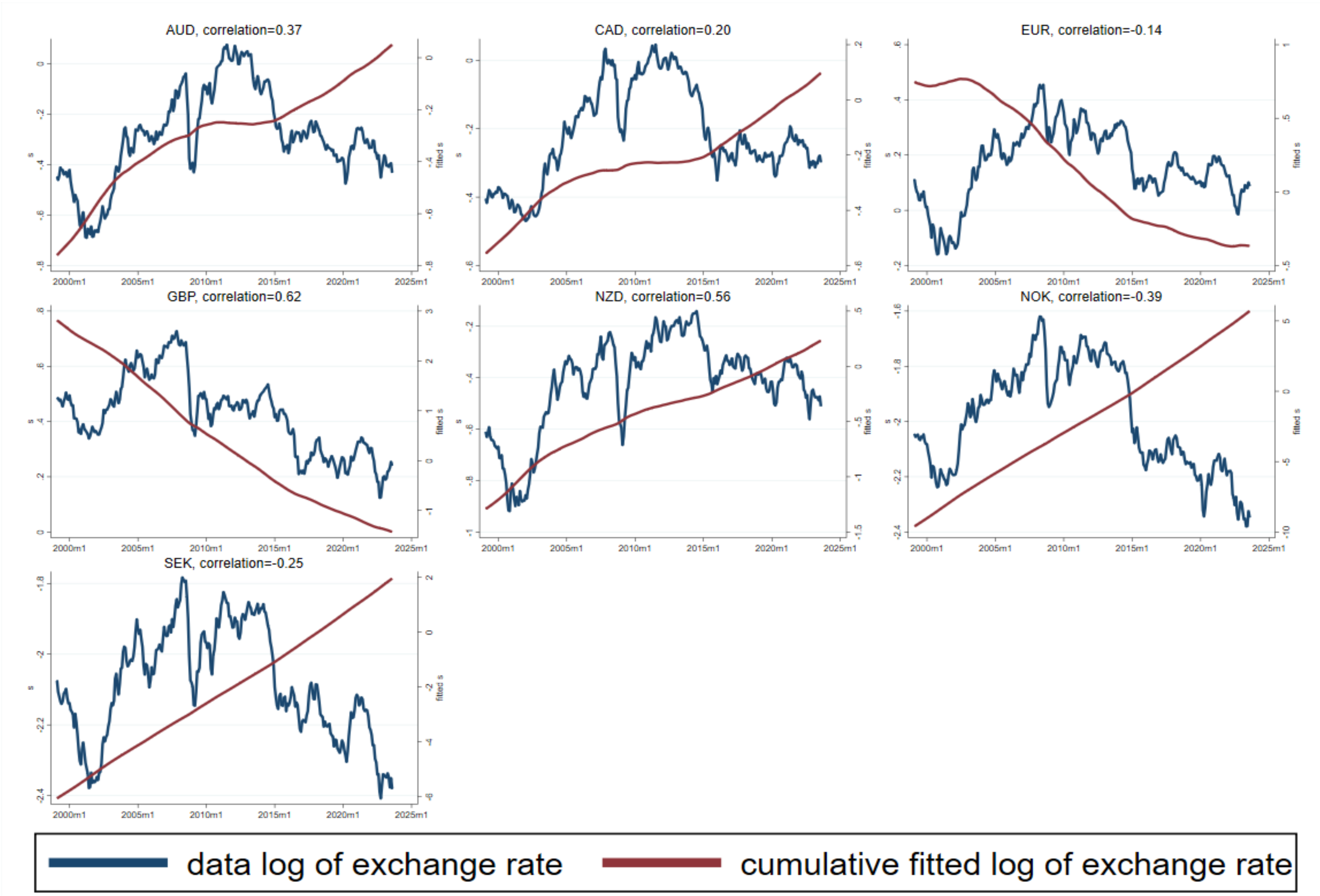
$$\Delta s_t = \alpha + \beta_5 \Delta RISK_t + \beta_7 \frac{TB}{GDP_t} + \beta_8 \eta_t + u_t$$



Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the data series. Correlations of the two series are reported in the subtitles. The exchange rate is defined as the U.S. dollar price of a foreign currency: an increase in value is a U.S. dollar depreciation.

Figure 1C: Fitted value generated from only lagged real exchange rate

$$\Delta s_t = \alpha + \beta_6 q_{t-1} + u_t$$



Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the data series. Correlations of the two series are reported in the subtitles. The exchange rate is defined as the U.S. dollar price of a foreign currency: an increase in value is a U.S. dollar depreciation.

Table C1: Baseline regression with only monetary variables

$$\Delta s_t = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_t^* + \beta_3 \pi_t + \beta_4 \pi_t^* + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta \eta_t + u_t$$

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
Δr_t	-1.00***	-1.57***	-2.19***	-1.19***	-0.90**	-1.71***	-1.68***	-1.38***	-1.38***
	(-2.91)	(-5.03)	(-7.30)	(-4.40)	(-2.56)	(-5.32)	(-5.27)	(-5.00)	(-4.99)
Δr_t^*	1.44***	1.16***	1.97***	1.84***	1.27***	0.29	0.66*	0.97***	0.97***
	(4.49)	(3.63)	(5.02)	(4.83)	(3.32)	(1.34)	(1.84)	(4.13)	(4.13)
π_t	-0.25**	-0.23*	-0.57***	-0.18*	-0.52***	-0.14	-0.42***	-0.30***	-0.30***
	(-2.19)	(-1.71)	(-4.15)	(-1.80)	(-3.81)	(-1.39)	(-3.62)	(-3.67)	(-3.71)
π_t^*	0.04	0.18	0.52***	0.10	0.32**	-0.18	0.21**	0.16**	0.15**
	(0.29)	(1.07)	(3.86)	(0.97)	(2.22)	(-1.41)	(2.30)	(2.07)	(2.10)
N	298	298	297	298	298	298	298	2085	2085
F	9.49	6.88	16.97	10.21	6.83	10.86	9.66	9.91	10.04
R2	0.11	0.09	0.19	0.12	0.09	0.13	0.12		0.09
R2_adj	0.10	0.07	0.18	0.11	0.07	0.12	0.10		
R2_within								0.09	

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the measure of the U.S. convenience yield relative to the foreign country, using 1-year government bond rates, as in Engel and Wu (2023).

Table C2: Baseline regression with only non-monetary variables

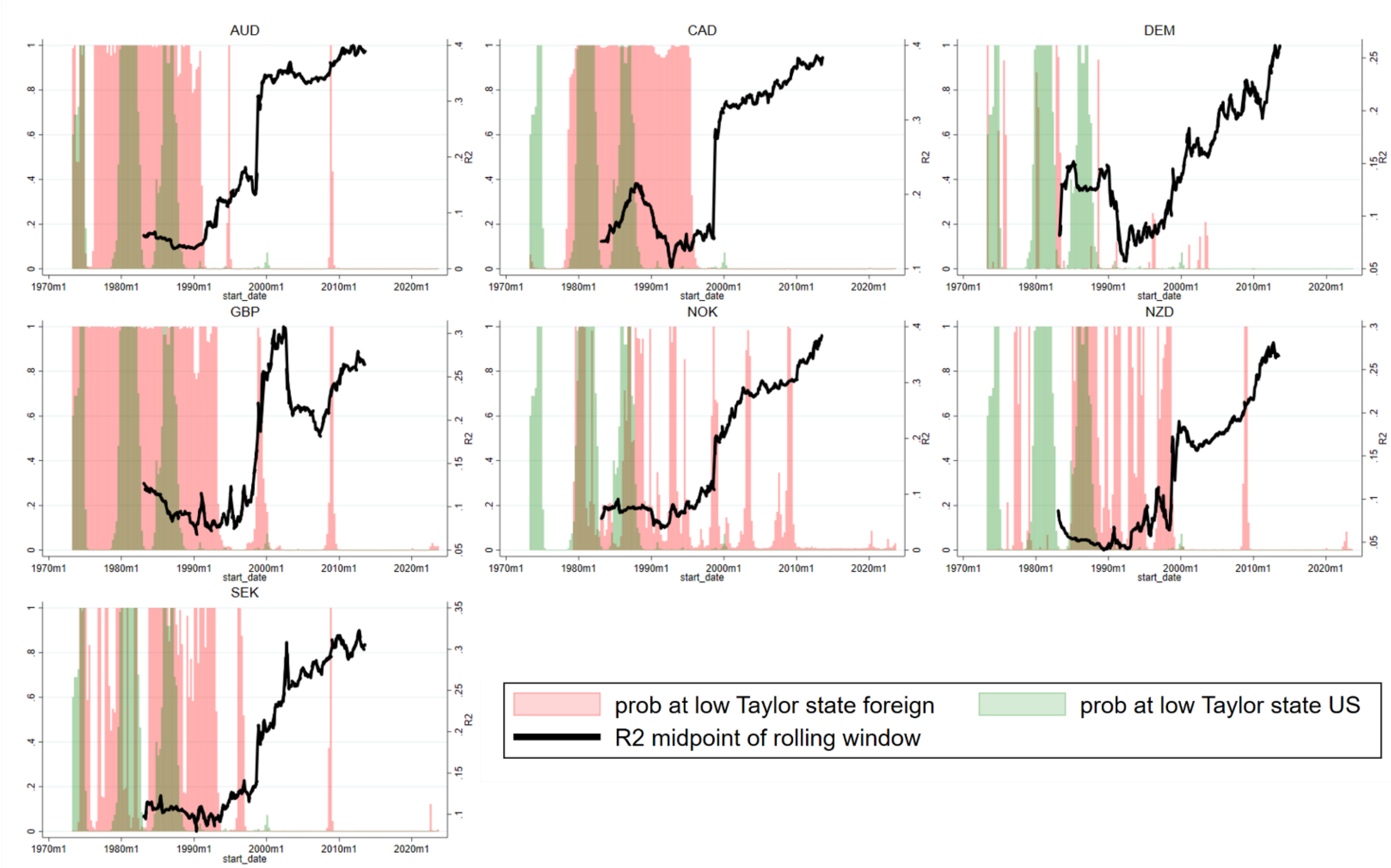
$$\Delta s_t = \alpha + \beta_5 \Delta RISK_t + \beta_6 q_{t-1} + \beta_7 \frac{TB}{GDP_t} + \beta_8 \Delta\eta_t + u_t$$

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect	Panel pooled
$\Delta RISK_t$	-0.04***	-0.02***	-0.01***	-0.02***	-0.03***	-0.02***	-0.02***	-0.02***	-0.02***
	(-10.45)	(-9.17)	(-3.10)	(-5.33)	(-7.09)	(-7.06)	(-5.81)	(-6.06)	(-5.98)
q_{t-1}	-0.01	-0.01	-0.01	-0.01	-0.01	-0.00	-0.00	-0.01	0.00
	(-1.17)	(-1.13)	(-1.50)	(-0.82)	(-1.29)	(-0.47)	(-0.19)	(-1.15)	(0.46)
$\frac{TB}{GDP_t}$	-0.27	-0.35***	-0.32*	-0.21	-0.18	-0.38**	-0.29	-0.29**	-0.26**
	(-1.60)	(-3.01)	(-1.94)	(-1.19)	(-1.00)	(-2.05)	(-1.54)	(-2.24)	(-2.01)
$\Delta\eta_t$	-1.42	-2.25***	-0.78	-1.56*	-1.60**	-1.58**	-0.31	-1.23*	-1.27**
	(-1.47)	(-2.65)	(-0.73)	(-1.69)	(-2.09)	(-2.14)	(-0.43)	(-1.93)	(-2.00)
N	296	296	295	296	296	296	296	2071	2071
F	29.34	28.19	3.73	8.74	15.30	15.38	9.30	11.24	10.45
R2	0.29	0.28	0.05	0.11	0.17	0.17	0.11		0.15
R2_adj	0.28	0.27	0.04	0.09	0.16	0.16	0.10		
R2_within								0.15	

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors. r_t and r_t^* are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign CPI inflation rate, $RISK_t$ is the first principal component of five risk variables. q_{t-1} is the real exchange rate in the previous period. TB/GDP_t is the trade balance to GDP of the U.S. η_t is the measure of the U.S. convenience yield relative to the foreign country, using 1-year government bond rates, as in Engel and Wu (2023).

Appendix D: Robustness of the Markov switching Taylor rule estimation

Figure D1 Markov switching Taylor rule estimation without probability constraint



Note: The figure reports the probability of $\phi_{\pi j}$<1 (the condition that Taylor principle fails) for the US in the shaded green region and the foreign country of the title of each subfigure in the shaded red region. The overlapped area is shaded with brown color. The coefficients of $\phi_{\pi j}$ at the high and low Markov states are reported in the subfigure title. The black lines report the R^2 measure of 20 year rolling window regressions in equation (1). For each date, the black line indicates the midpoint of the rolling window.