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

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Abstract

Empirical exchange-rate models fit the U.S. dollar very well in the 21st century. A "standard" model that includes real interest rates, a measure of expected inflation for the U.S. and the foreign country, and the U.S. comprehensive trade balance—augmented with measures of global risk and liquidity demand—is well supported by the data. From the 1970s to the 1990s, the fit of the model was poor, but it has improved for both traditional and risk variables almost monotonically to the present day. We provide evidence that better monetary policy has led to this improvement.

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1. Introduction

In the late 1970s and early 1980s, the "asset market" approach to exchange rate determination emerged that posited three important determinants of exchange rates: First, monetary models (such as Dornbusch (1976) and Frankel (1979)) highlighted the role of real interest rates. An increase in a country's real interest rate leads to appreciation. Second, the asset market approach emphasized that exchange rates 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. Third, 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.

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." Obstfeld and Rogoff (2000) coined "the exchange-rate disconnect puzzle" as "a name that alludes broadly to the exceedingly weak relationship between the exchange rate and virtually any macroeconomic aggregates"

Since 2000, however, 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 (Lilley et al 2022, Jiang et al 2021, Engel and Wu 2023).

We make two contributions to the literature by providing an integrated and comprehensive empirical investigation of exchange rates models. First, the paper documents a little-recognized phenomenon: that in the 21st century, a *conventional* empirical exchange rate model explains the data well with expected signs consistent with theory predictions. Both the traditional variables motivated from the 1970s and measures of global risk and liquidity motivated from the recent literature help account for dollar exchange rates and contributed equally to the good fit.[1] The empirical model we fit for the 21st century is precisely of the form that led economists to believe

[1] We include two measures of financial conditions (the aforementioned global risk and liquidity variables) as explanatory variables that were not important in the earlier studies, but, as the literature has noted, have become more important in recent decades. But, as we highlight, the "traditional" explanatory variables such as interest rates, are no longer as disconnected from the exchange rate, as they were previously.

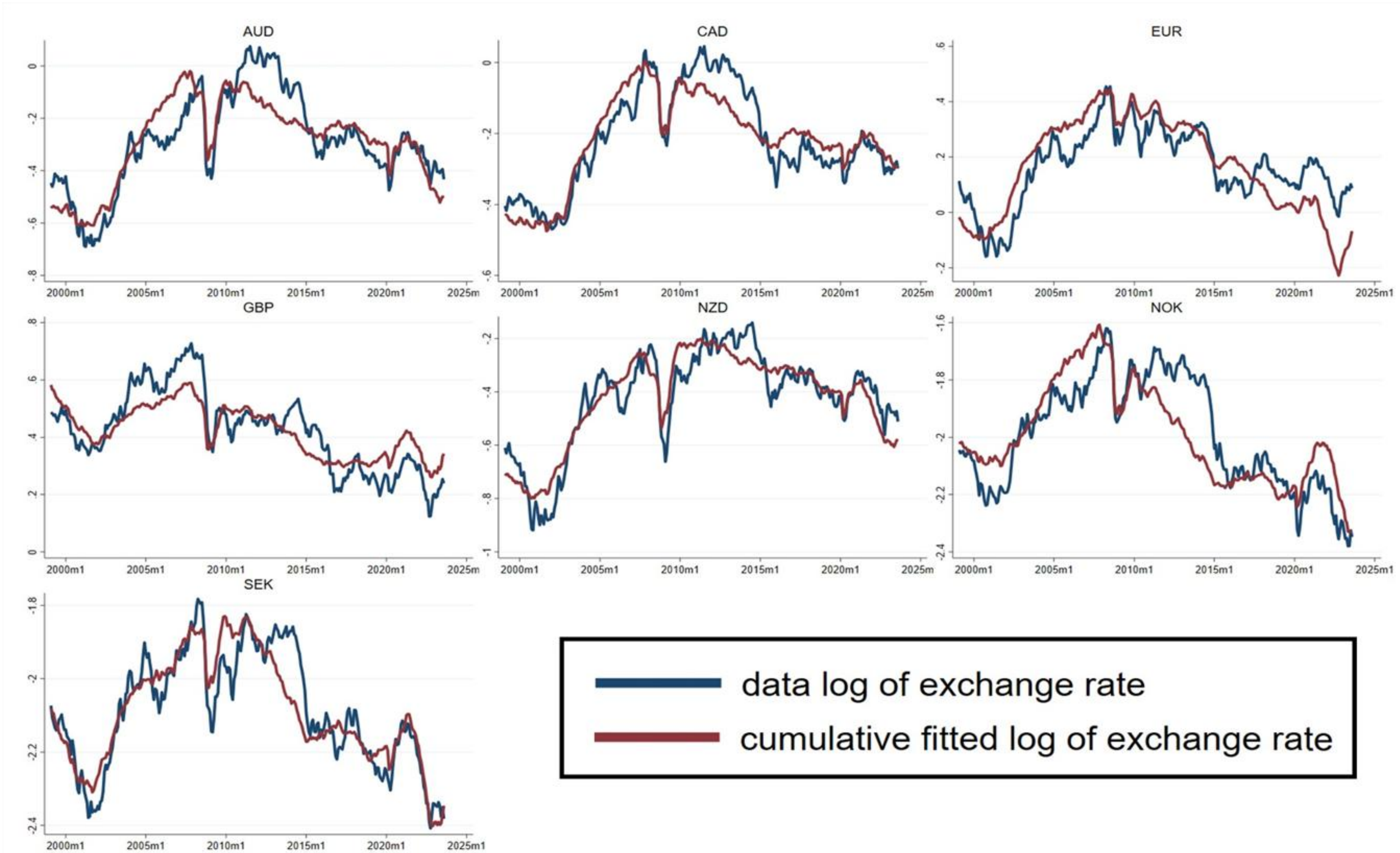
previously – based on 20th century data – that the exchange rate was disassociated from its supposed fundamental determinants. This disconnect is no longer so apparent in the data.

Second, we offer an explanation from the disconnect to the recent good fit. It is not enough to have a model that fits well on recent data – we need an explanation for why the model did not work previously. We hypothesize that inflation credibility among advanced economies has been substantially improved in the 1990s. That leads to a more predictable and transparent central bank policy rule, which subsequently strengthened the link between monetary policy stances and exchange rates. We provide various types of causal evidence to establish the relationship from central bank credibility to improvement of regression fit of the exchange rate model.

Figure 1 demonstrates how well the model can track the exchange rate dynamics observed in the data. In each subfigure, the blue line is the log of the exchange rate, and the red line is the cumulative fitted value of the change in the log of the exchange rate from the model described in Section 2. Evidently, the fitted series captures the exchange rate dynamics well. Taking the euro 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.

These estimates are on monthly changes in log exchange rates. The model is not fit on levels, where high correlation can arise by coincidence when exchange rates and the explanatory variables are highly persistent. Another important point to note is that the good fit of the model does not come from, in essence, regressing the exchange rate on the exchange rate. That is, the exchange rate itself is not used in the construction of any of the regressors, so the good fit is not artificial. Monthly changes are the horizon over which Meese and Rogoff, Cheung et al. (2005), and Rossi (2013) found previously that models did not fit. We find the R^2 for monthly changes in the baseline regression is in the range of 0.25 to 0.30 for the dollar relative to the euro, the pound, the Swedish krona, Norwegian krone, and New Zealand dollar, and 0.40 for the Canadian and Australian dollars. The fit on monthly data is even tighter when we allow for slightly richer dynamics, and on quarterly data, where some higher frequency shocks are smoothed out, the fit of the model is extremely strong. (Figure A1 in Appendix B shows the same figure at a quarterly frequency. For example, the R^2 for the dollar/euro rate is 0.54, and the within R^2 for the panel regression is 0.55.)

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. 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 pre-2000 sample (the "old days"), the fit was poor – the variables are usually statistically insignificant and sometimes have the wrong sign; and the R^2 values are low. But there is a near-monotonic increase in 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 of out-of-sample 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, but the Clark-West statistic strongly rejects the null of no explanatory power relative to the random walk in later periods.

Through a comprehensive battery of robustness checks, we emphasize that the strong explanatory power of these exchange rate models emerged well before 2008 and is not merely an artifact of outlier volatility from the 2008–2009 financial crisis. We demonstrate that traditional macroeconomic fundamentals play a substantial role in driving the improved performance of the model whether or not the period 2008-2009 is included in the sample.

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 regimes may explain this. Before countries either explicitly or implicitly adopted inflation targeting, the Taylor principle was not satisfied, which weakens the link between interest rates, expected inflation, and exchange rates. We contend that as credibility increased, this phenomenon decreased, and the fit of the standard model improved. 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, 2001) 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 before 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 poor explanatory power of standard models should persist into the 1990s by this line of reasoning. We show that the improved fit of the model over time corresponds well to a measure of central bank independence that is increasing in the late 20th century.

Ilzetzki 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 model of the link.

We show that the poor fit of the model in the early part of the sample arises 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's real interest rates usually did not contribute in the way theoretical 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.

In the second part of the paper, we strengthen our analysis by providing causal evidence on how improved central bank credibility enhances the fit of exchange rate regressions. We do so through three complementary approaches: a staggered difference-in-differences (DiD) analysis, an instrumental variable (IV) regression, and a high-frequency identification strategy.

We estimate Taylor rules for each country, allowing for a regime change using a switching model with a one-time, endogenously determined breakpoint. Across all countries, the early part of the sample aligns with a regime in which the Taylor principle is not satisfied, while the later period reflects a regime where it is. Using these estimated inflation credibility cutoff dates, we conduct two exercises. First, we show that the relationship between interest rates and exchange rates changes substantially before and after the credibility threshold. Second, we exploit these cutoff dates to implement a staggered DiD regression. The results demonstrate that the "treatment" of establishing inflation credibility leads to a significantly stronger relationship between exchange rates and theory-implied macroeconomic fundamentals.

In the IV analysis, we exploit the insight that countries often undertake policy reforms in response to successful policy implementations in economically or regionally connected peers (Acemoglu et al. (2019)). Specifically, we instrument the central bank credibility index from Romelli (2024) using the average index measured in economically or geographically connected countries. This instrument enables us to isolate improvements in central bank credibility that are plausibly exogenous to a country's own macroeconomic conditions. We find that increases in a country's central bank credibility index are associated with a significantly better fit of the exchange rate regression model.

Finally, we examine how exchange rates and interest rates respond to inflation news in a high-frequency setting. When monetary policy is credible and agents expect the central bank to target inflation, the currency should appreciate following a positive inflation surprise (as in Andersen et al., 2003; Clarida and Waldman, 2008). We find that when inflation announcements exceed survey expectations, markets immediately respond with an upward revision in expected future policy interest rates and an appreciation of the currency. These findings provide direct evidence of a strong interest rate–exchange rate linkage when the central bank is credibly committed to inflation targeting.

Literature Review

Our work ties into four 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.

Dornbusch's (1976) classic presentation of the overshooting model 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.

In forward-looking models, exchange rates are connected not just to current economic fundamentals, but, as Frenkel (1981) emphasizes, to news about the future. Some studies, such as Engel and Frankel (1984), Andersen et. al. (2003), and Clarida and Waldman (2008), examine how announcements of measures of economic aggregates affect the exchange rate. 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. Stavrakeva and Tang (2024) link expectations from surveys of traders to exchange-rate movements.

As investors are required to hold more of a country's debt in their portfolio of risky assets, they require a higher expected return. The portfolio balance model of Kouri (1976, 1981), Dooley and Isard (1982), and Hooper and Morton (1982), predicts as a country increases its debt, the currency initially depreciates in response to an increase in debt to generate an expected appreciation. Prominent recent revivals of the portfolio balance model include Gabaix and Maggiori (2015), Della Corte et al. (2016), Fang and Liu (2021), Camanho et al (2022), 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 large recent literature. The central role of the U.S. and its status as a safe-haven currency have been explored by, among others, Maggiori (2017), Farhi and Maggiori (2018), Rey (2015, 2016), Miranda-Agrippino and Rey (2020), Gourinchas and Rey (2022), Kekre and Lenel (2024a) and Akinci et al (2022). Lilley et al. (2022) and Obstfeld and Zhou (2023) find strong empirical support for the relationship between the value of the dollar and measures of financial fragility. Papers have linked liquidity returns, or convenience yields, to exchange rates, including Avdjiev et al (2019), Du et al. (2018), Krishnamurthy and Lustig (2019), Jiang et al. (2020, 2021, 2024), and Engel and Wu (2023).

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 (2021, 2025). There are two sides to the disconnect puzzle. First, the models that are supposed to explain exchange rates did not work. Second, exchange rate changes seem to have little impact on economic variables. 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.[2]

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

[2] See Kollmann (2001), Jeanne and Rose (2002), Devereux and Engel (2002), and Itskhoki and Mukhin (2021, 2025).

(2004), Coibion and Gorodnichenko (2011), Lubik and Schorfheide (2004), Bhattarai et al. (2016), Molodtsova et al (2011) and Hirose et al. (2020). 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.

In contemporaneous work, Kekre and Lenel (2024b) build a general equilibrium model that reproduces many of the features of the empirical model we estimate here. 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.

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. 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 leads to a breakdown in the empirical relationship between exchange rates and the fundamentals and then present evidence to support the claim that the improved fit of the models coincides with stricter inflation targeting.

2. Empirical exchange rate models fit the data

In this section, we present the empirical model 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 the 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

We estimate a quite standard empirical model for exchange-rate determination, augmenting the fundamentals introduced in the 1970s and 1980s (as in Meese and Rogoff (1983)) with the global risk and liquidity variables the more recent literature has emphasized. We estimate the following monthly regression from January 1999 to August 2023:

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

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 describe heuristically here how each variable affects exchange rates, following the standard discussion of these empirical relationships in the literature.[3] There are economic forces that might change the signs of each of these relationships, but at the end of section 3a we argue that the relationships here are plausible. Appendix E shows how this equation can arise in a standard New Keynesian model.

1. *Real interest rates*. An increase in the home real interest rate leads to an appreciation 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. As a measure of expected inflation in the U.S. and all the foreign countries used to construct the real interest rates, we take the inflation

[3] See Meese and Rogoff (1983), Frankel (1979), Frenkel (1981), Hooper and Morton (1982), and more recently, Obstfeld and Zhou (2023).

rate over the previous year. This has the benefit of being a consistent, easily reproducible measure for all countries.[4]

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 inflation over the past year for the U.S. and the foreign country in the regression. 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.

   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. 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 is associated with a dollar appreciation.

   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.

[4] Atkeson and Ohanian (2001) find that the inflation rate over the past year beats a range of other potential forecast variables in terms of mean-squared forecast error.

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 our measure 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.

The inclusion of the trade balance variable allows us to follow exactly Meese and Rogoff's (1983) empirical model. That study justifies the inclusion of the trade balance, as we do, based on the empirical portfolio balance model of Hooper and Morton (1982). Depending on the source of the shocks to the economy, the sign of the coefficient could be positive or negative, but Meese and Rogoff rely on the portfolio balance effect to be dominant, and our simulation of a simple New Keynesian model in Appendix E shows that interpretation is defensible.

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 falls 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.

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.

In section 4 and Appendix E, we show that the regression predictions can be derived from a standard two-country New Keynesian model. Specifically, the model's specification of household and firm behavior is typical, but all financial trades must be intermediated through a risk-averse global arbitrageur (as in Itskhoki and Mukhin, 2021, Kekre and Lenel 2024b and Cormun and De Leo 2026.) But we stress that we are not presenting a test of that particular model. Our point is that the very regressions that led the profession to believe monetary sticky-price models were irrelevant for understanding the exchange-rate data, when augmented with measures of global risk and liquidity, no longer exhibit the disconnect that they did in the 20th century.

b. Data

Our analysis focuses on two sample periods, January 1999-August 2023 and March 1973-August 2023. We use monthly average exchange rate data to smooth out shorter-term movements. We obtain nominal exchange rate data from Federal Reserve Board of Governors H.10 release. Home and foreign consumer price indexes 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. Nominal government bond interest rates 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 go 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 rates. Second, the liquidity ratio in Bianchi et al (2021), which is constructed from FRED data as the sum of U.S. dollar financial commercial paper (DTBSPCKFM) and short-term funding to U.S. banks is demand deposits (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.) 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. Supplementary Appendix Table A1 provides the detailed data information 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. The last column reports the panel regression of all seven currencies with fixed effects. The R^2 and F-statistics of these regressions are high. The R^2s range from 0.22 for GBP to 0.36 for AUD. The F-statistics are all above 10, which corresponds to a p-value smaller than 0.0001 and indicates a very high joint significance of the variables in explaining exchange rates. Figure 1, discussed in the introduction, shows the cumulative fitted values from these regressions.

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 -0.25 for U.S. inflation, which indicates a one percentage point increase in year-on-year inflation is associated with 0.25% 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 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 TB/GDP are negatively significant which 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 indicates if the TB/GDP improves by one percentage point, the U.S. dollar immediately appreciates by 0.54%. Overall, these traditional variables that are emphasized in Meese and Rogoff (1983) are all statistically significant.

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. Appendix Table A7 presents the regressions with the liquidity ratio measure from Bianchi et al. (2021) replacing the relative convenience yield. The performance of the model is quite similar.

Finally, we find some evidence of long run PPP, 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.

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

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

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel (FE)
Δr_t	-1.12***	-1.66***	-2.34***	-1.37***	-0.92***	-1.85***	-1.90***	-1.47***
	(-3.91)	(-5.85)	(-8.06)	(-4.79)	(-2.98)	(-5.56)	(-5.45)	(-5.83)
Δr_t^*	1.12***	1.16***	2.21***	1.80***	1.06**	0.23	0.79*	0.92***
	(3.59)	(4.22)	(6.01)	(3.59)	(2.34)	(1.02)	(1.66)	(3.82)
π_t	-0.25***	-0.21	-0.69***	-0.33***	-0.45***	-0.21**	-0.57***	-0.34***
	(-2.76)	(-1.50)	(-4.58)	(-2.61)	(-4.13)	(-2.03)	(-5.66)	(-5.24)
π_t^*	0.03	0.14	0.53***	0.14	0.24**	-0.19	0.25***	0.15**
	(0.26)	(0.90)	(3.84)	(1.39)	(2.02)	(-1.49)	(2.94)	(2.56)
$\frac{TB}{GDP_t}$	-0.48***	-0.45***	-0.63***	-0.73***	-0.37*	-0.66***	-0.80***	-0.54***
	(-2.74)	(-4.36)	(-4.24)	(-3.30)	(-1.88)	(-3.38)	(-4.03)	(-4.35)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01***	-0.01***	-0.02***	-0.02***	-0.02***	-0.02***
	(-9.12)	(-12.81)	(-2.97)	(-4.38)	(-7.24)	(-5.74)	(-5.04)	(-8.42)
$\Delta \eta_t$	-1.92**	-2.33***	-0.86	-1.52**	-1.56**	-1.20*	-0.69	-1.38***
	(-2.24)	(-4.22)	(-0.70)	(-2.15)	(-2.06)	(-1.95)	(-0.78)	(-2.92)
q_{t-1}	-0.01	-0.01	-0.02**	-0.03***	-0.01	-0.03***	-0.01	-0.01**
	(-1.26)	(-1.63)	(-2.43)	(-3.22)	(-1.52)	(-2.69)	(-1.43)	(-2.12)
N	296	296	295	296	296	296	296	2071
F	26	40	17	18	28	28	18	35
Adjusted R^2	0.36	0.36	0.25	0.22	0.22	0.30	0.24	0.25(within)
Shapley (Traditional)	29%	30%	85%	59%	34%	46%	58%	43%
Shapley (Financial)	69%	69%	11%	36%	64%	50%	40%	55%

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. Standard errors are Newey West standard errors with 12 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 12 lags. 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). Shapley (traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$). Shapley (financial) reports the same Shapley value decomposition for $\Delta RISK_t$ and $\Delta \eta_t$

What are the relative contributions of the traditional versus the more recent macro-financial variables (risk and convenience yield)? The last two rows of Table 1 report the Shapley value regression decomposition for these two sets of variables.[5] The traditional variables are those studied in Meese and Rogoff (1983), including the change of real interest rates, inflation, and the trade balance. The financial variables are the change of risk variable and the change of convenience yield measure. The Shapley value of a regression calculates the marginal contribution of a set of variables across every possible combination of the other regressors. Therefore, it summarizes each variable's average contribution to the overall explanatory power of the regression. The Shapley values of all variables in the regression sum to 100% of the regression model's total explained variation (R^2). The Shapley value has an advantage of calculating a variable's contribution while still controlling for other variables.

The second last row reports that the Shapley values of the traditional variables range from 29% to 85%, with an average of 49% and a panel regression Shapley value of 43%. This indicates that the traditional variables account for roughly half of the explained variation. The last row reports that financial variables account for roughly the same, with an average of 48% for the seven currencies and a Shapley value of 55% for panel fixed effect regression.[6]

[5] See Kruskal (1987), Israeli (2007), Huettner, F. and M. Sunder (2012) on the foundations of the Shapley value as a measure of relative contribution of a variable to explanatory power in a regression. Cox, et al. (2024), Baumann et al (2024), and Crouzet et al (2026) are some recent examples that use the Shapley value in macro context.
[6] The sum of traditional variable Shapley contribution and financial variable Shapley contribution does not add up to 100% since there are other variables, including the lagged value of real exchange rate and the fixed effects.

Table 2: Regression with 10-year interest rates and lagged variables

LHS: Δs_t	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel (FE)	Panel (quarterly)
$\Delta r_t^{10y,US}$	-8.92***	-5.27***	-6.40***	-1.48	-7.23***	-5.03***	-5.90***	-5.63***	-7.51***
	(-9.05)	(-3.98)	(-6.37)	(-1.25)	(-5.83)	(-5.21)	(-6.59)	(-9.25)	(-6.16)
$\Delta r_t^{10y,Foreign}$	8.74***	6.90***	6.46***	1.94	6.25***	5.71***	5.94***	5.91***	8.21***
	(7.54)	(4.24)	(5.28)	(1.19)	(4.35)	(5.07)	(4.23)	(6.11)	(6.32)
π_t^{US}	-8.49***	-4.33***	-4.86***	-0.76	-7.17***	-3.57***	-4.86***	-4.75***	-6.84***
	(-7.85)	(-3.34)	(-4.29)	(-0.63)	(-5.34)	(-3.42)	(-4.98)	(-7.56)	(-5.92)
$\pi_t^{Foreign}$	8.10***	6.31***	5.01***	1.20	6.13***	5.46***	5.95***	5.50***	7.55***
	(7.06)	(4.06)	(3.46)	(0.71)	(4.35)	(4.38)	(4.44)	(5.85)	(5.73)
$\frac{TB}{GDP_t}$	0.03	-0.33**	-0.56***	-0.42*	-0.25	-0.54**	-0.59***	-0.39***	-1.10***
	(0.19)	(-2.56)	(-3.12)	(-1.80)	(-1.31)	(-2.55)	(-2.90)	(-3.14)	(-3.40)
$\Delta Risk_t$	-0.04***	-0.02***	-0.01***	-0.01***	-0.03***	-0.02***	-0.02***	-0.02***	-0.03***
	(-11.92)	(-8.03)	(-2.94)	(-4.08)	(-6.94)	(-4.85)	(-5.18)	(-8.38)	(-8.37)
$\Delta \eta_t$	-1.37*	-2.58***	-1.08	-2.06***	-2.22***	-1.47**	-1.11*	-1.53***	-2.86***
	(-1.65)	(-3.37)	(-0.96)	(-3.10)	(-3.13)	(-2.17)	(-1.83)	(-3.35)	(-3.18)
q_{t-1}	-0.06***	-0.00	-0.03***	-0.05***	-0.06***	-0.02**	-0.03***	-0.02***	-0.04***
Lagged variables below	(-4.93)	(-0.41)	(-3.00)	(-3.41)	(-4.08)	(-2.39)	(-2.98)	(-3.92)	(-2.92)
$r_{t-1}^{10y,US}$	-2.19***	-0.31	-0.50	-0.32	-1.83***	-0.77**	-0.66**	-0.58***	-1.96***
	(-4.10)	(-1.02)	(-1.53)	(-0.71)	(-3.32)	(-2.17)	(-2.33)	(-3.95)	(-5.19)
$r_{t-1}^{10y,Foreign}$	1.67***	0.38	0.33	0.57	1.16***	0.81***	0.68***	0.53***	1.70***
	(3.87)	(1.46)	(1.27)	(1.25)	(3.03)	(2.90)	(2.78)	(4.06)	(5.11)
π_{t-1}^{US}	6.19***	3.87***	3.84***	0.06	5.01***	2.49**	3.68***	3.87***	4.15***
	(5.17)	(2.77)	(3.38)	(0.05)	(3.58)	(2.30)	(3.58)	(6.11)	(3.50)
$\pi_{t-1}^{Foreign}$	-6.58***	-5.89***	-4.36***	-0.71	-4.97***	-4.73***	-5.16***	-4.94***	-5.80***
	(-5.12)	(-3.58)	(-3.06)	(-0.42)	(-3.49)	(-3.56)	(-3.81)	(-5.13)	(-4.38)
$Risk_{t-1}$	-0.00***	-0.00	-0.00	-0.00*	-0.00**	-0.00	-0.00	-0.00*	-0.00**
	(-2.63)	(-0.94)	(-0.02)	(-1.82)	(-2.35)	(-0.43)	(-1.50)	(-1.82)	(-2.17)
η_{t-1}	-0.35	-0.76	-1.74***	-1.86***	-0.91	-0.95*	-1.30***	-0.89***	-1.50
	(-0.48)	(-1.26)	(-2.92)	(-2.81)	(-1.63)	(-1.80)	(-3.26)	(-2.93)	(-1.65)
N	296	296	295	296	296	296	296	2071	692
F	65	31	20	16	30	30	35	60	48
Adjusted R^2	0.48	0.40	0.32	0.24	0.31	0.37	0.35	0.33 (within)	0.55 (within)
Shapley (Traditional)	46%	41%	80%	47%	37%	60%	64%	53%	56%
Shapley (Financial)	52%	59%	16%	48%	57%	38%	34%	45%	44%

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. Standard errors are Newey West standard errors with 12 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 12 lags for monthly regression and 4 lags for quarterly regression. Shapley (traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$ and its lags). Shapley (financial) reports the Shapley value decomposition for $\Delta RISK_t$ and $\Delta \eta_t$ and its lags.

Table 2 makes three modifications. First, it includes lagged levels of the right-hand-side variables, except the trade balance and the lagged real exchange rate. The real interest rate, inflation, and financial variables are persistent, but likely stationary, so including lagged levels protects against over-differencing and essentially just allows for different coefficients for the current and lagged variables. Second, instead of using 3-month interest rates as in our baseline regression, we use the 10-year government rates, which presumably are a better measure of long-term monetary policy stance. These are measured as yields to maturity, taken from the OECD Long-term Rates Database. Finally, we also report the panel fixed effects regression at a quarterly frequency, which is more compatible with standard open economy theoretical models (the individual country regressions are reported in Appendix Table A5).

Three noticeable observations arise. First, after controlling for lags, the fit of the model is good, with all expected signs and t-statistics comparable to the baseline regression. The adjusted R^2s are increased by about 8 percentage points to an average of 35%. Second, the Shapley contribution of traditional variables has generally increased to an average of 54%, highlighting the importance of these variables in explaining exchange rate movements. Finally, the quarterly panel regression reports an impressive within R^2 of 55% (The R^2 values for the individual currency quarterly regressions range from 0.45 to 0.72.) These findings offer compelling empirical evidence that macroeconomic fundamentals are strongly connected with exchange rate dynamics. Therefore, treating the exchange rate disconnect as a stylized fact or a primary modeling objective *may no longer be empirically justified*, and instead the emphasis should be on the relatively strong relationship between exchange rates and the macro and financial variables.

In Appendix B, we illustrate that the good fit of Table 2 is not just a U.S. dollar phenomenon. We report panel regressions that are similar to the panel column of Table 2 but treating each foreign currency as the base currency and excluding the U.S. dollar from the sample. Because trade balance data is not available at a monthly frequency and a measure of risk appetite is not readily available for other countries, we drop these two variables and leave the (home minus foreign) relative convenience yield as the only financial variable. Appendix Table A9 shows that the real interest rate, inflation, lagged real exchange rate, and convenience yield variables are all highly significant with the expected signs. The R^2 of these regressions using foreign currencies as the base currency is generally lower at about 10% to 15%, but comparable to the US case (18%) when dropping the trade balance and the global risk measure. On the other hand, the Shapley contributions of the traditional variables are higher, ranging from 78% to 96%.

In conclusion, the empirical evidence presented above suggests a reconsideration of the traditional exchange rate disconnect puzzle by highlighting three key findings. First, the model achieves a solid fit, with variables generally yielding high joint significance and coefficients that align with theoretical expected signs. Second, alongside modern financial factors like risk and convenience yields, the explanatory contribution of traditional macroeconomic variables—namely real interest rates, inflation, and trade balances—remains substantial, frequently accounting for over half of the model's explanatory power. Finally, this framework appears reasonably robust, as it performs well across various measures and maintains its explanatory relevance when applied to various other non-U.S. base currencies. Together, these results indicate that macroeconomic fundamentals play a meaningful role in shaping exchange rate movements.

The literature has interpreted such regressions as causal, so it is worth taking a moment to consider why. 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 improbable 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. When the currency was depreciating, the policymaker would increase interest rates, but we find the opposite partial correlation.

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 if exchange rates passed through significantly to consumer prices. 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.

Some open-economy macroeconomic models predict that a depreciation will lead to an increase in the trade balance as households and firms switch expenditures away from foreign goods. To the extent this channel is effective, it would tend to offset the portfolio balance effect our model posits, which relates trade deficits of the U.S. to a depreciating currency.

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.

The important point is that the regression results that led the profession to move away from sticky-price monetary models no longer send that message. Here we are simply explaining why these regressions are generally interpreted as demonstrating some sort of causality. The objective of fully-specified models of exchange-rate behavior might better be aimed at accounting for these correlations rather than models that find little connection as there was in the past.[7]

b. Model estimation from 1973

We have shown that the exchange-rate model is successful for the sample from 1999-2023. In this subsection, we extend the sample back 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, we use 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 in pre-1999, we drop the variable for this exercise.

In-sample analysis

We first provide a snapshot of the pre-1999 regression results in Table 3, which reports the same regression as in Table 1 but changing the sample to Mar. 1973 to Dec. 1998. Table 3 shows a drastically different pattern. Most of the regression coefficients are not statistically significant, and many are opposite of the theoretical prediction. For example, the foreign interest rate coefficients for CAD, DEM, GBP are negative. The adjusted R^2 is less than 10% for most currencies. The panel regression R^2 is very low at 3%.

[7] As noted above, in section 4, we generate artificial data from a New Keynesian model laid out in detail in Appendix E. We estimate the exchange-rate regression with the artificial data and show that it can replicate the results with the actual data, when the Taylor condition is satisfied.

. Table 3: Baseline regression in the "old days", sample period Mar 1973-Dec 1998

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

	AUD	CAD	DEM	GBP	NZD	NOK	SEK	Panel (FE)
Δr_t	-0.72***	-0.03	-0.77***	-0.72***	-0.18	-0.65***	-0.86***	-0.66***
	(-2.83)	(-0.32)	(-4.02)	(-2.85)	(-0.36)	(-2.89)	(-4.03)	(-4.01)
Δr_t^*	0.03	-0.51***	-0.22	-0.62**	0.11	0.16	0.35***	0.02
	(0.14)	(-6.04)	(-0.46)	(-2.10)	(0.55)	(0.69)	(2.82)	(0.15)
π_t	0.08	-0.03	-0.09	0.07	-0.20*	0.07	0.02	0.00
	(1.25)	(-0.77)	(-0.90)	(0.73)	(-1.75)	(1.05)	(0.26)	(0.05)
π_t^*	0.02	0.05	0.26	-0.02	0.12	-0.06	-0.01	0.01
	(0.34)	(1.23)	(1.00)	(-0.36)	(1.49)	(-0.98)	(-0.26)	(0.47)
$\frac{TB}{GDP_t}$	0.19	-0.08	-0.77*	-0.53**	0.01	-0.17	-0.42*	-0.26
	(0.70)	(-0.60)	(-1.73)	(-2.23)	(0.03)	(-0.66)	(-1.85)	(-1.55)
$\Delta RISK_t$	-0.01**	-0.00	0.00	-0.00	0.00	-0.00	-0.00	-0.00
	(-2.57)	(-0.37)	(0.72)	(-0.68)	(0.54)	(-0.07)	(-0.88)	(-0.65)
q_{t-1}	-0.03*	0.00	0.01	-0.01	0.01	-0.01	0.00	-0.00
	(-1.96)	(0.08)	(0.45)	(-0.37)	(0.98)	(-0.67)	(0.38)	(-0.14)
N	311	311	311	311	205	311	311	1760
F	1.73	8.70	6.74	4.45	1.59	5.40	7.90	6.00
Adjusted R^2	0.03	0.12	0.05	0.08	0.01	0.02	0.06	0.03 (within)

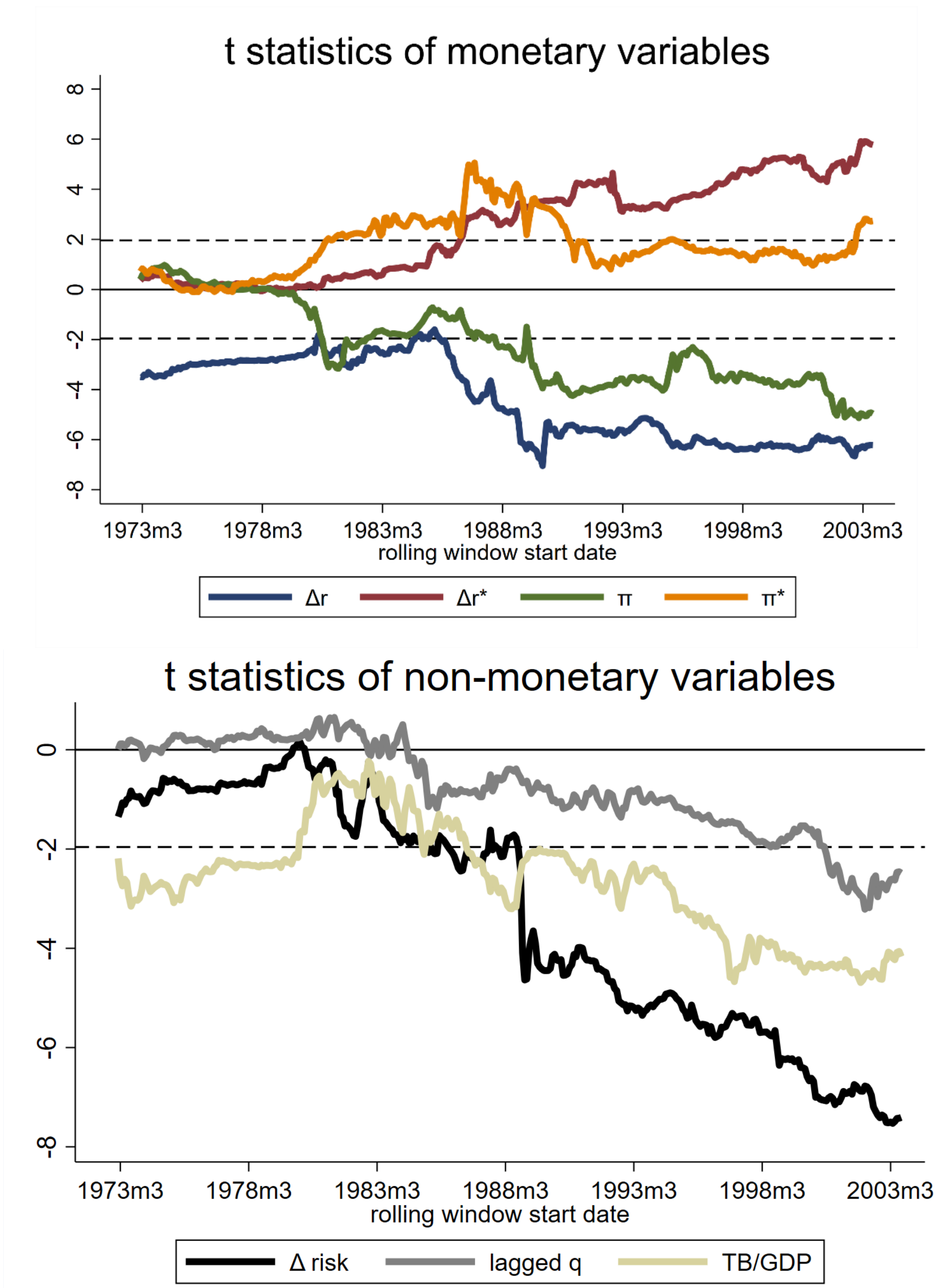
Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Mar 1973 to Dec 1998. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. Standard errors are Newey West standard errors with 12 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 12 lags. 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

It is instructive to look at the evolution of *t*-statistics to see when the empirical model contradicts theoretical predictions and when it begins to behave well. Figure 2 plots the t-statistics for each variable in the 20-year rolling-window panel fixed-effects regressions (country by country figures are reported in Appendix B Figure A3). The x-axis represents the start date of the rolling window. For example, a start date of March 1973 corresponds to the t-statistics from a regression using the sample period from March 1973 to February 1993. The final regression in the series starts in September 2003 and ends in August 2023. In the upper panel, we plot the *t*-statistics of the monetary variables: the change in the home real interest rate (blue line), the change in the foreign real interest rate (red line), the home inflation rate (green line), and the foreign inflation rate (orange line). In the lower panel, we plot the *t*-statistics for the change in the global risk measure (black line), the lagged real exchange rate (grey line), and the U.S. trade-balance-to-GDP ratio (gold line).

It is clear from Figure 2 that the model's relationships became "well behaved" in the second half of the sample. In the upper panel, the home real interest rate and home inflation rate become negatively significant starting around 1980 (which corresponds to the 1980–2000 sample period). Furthermore, these *t*-statistics become increasingly significant in the latter part of the sample. Similarly, the foreign real interest rate and foreign inflation rate become positively significant starting around 1986 (which corresponds to the 1986–2006 sample period). Therefore, there is strong evidence that these monetary variables become increasingly important over time.

In the lower panel of Figure 2, the *t*-statistics for the global risk measure, the lagged real exchange rate, and the trade-balance-to-GDP ratio all shift into negative territory (as the model implies) during the second half of the sample. However, these patterns do not hold in the first half. For regressions with start dates between 1973 and 1984, the t-statistics for all three measures are largely insignificant. The t-statistics for the global risk measure and the trade-balance-to-GDP ratio become marginally significant for rolling windows starting between 1985 and 1988. The global risk measure then experiences a sharp increase in statistical significance for start dates around 1988, which coincides with the inclusion of the 2008 global financial crisis in the 20-year rolling window.

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. Dashed lines indicate +1.96 and -1.96, critical values for 5% significance. X-axis corresponds 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.

To evaluate how much these factors collectively improve the explanatory power of exchange rate models, in Figure 3, we report the adjusted R^2 (blue line) for the same panel fixed-effects 20-year rolling regression shown in Figure 2.[8] We also report the corresponding R^2 contributions of the traditional (red line, including the change of real interest rates, inflation, and the trade balance) and financial (green line, including the change of risk measure) variables, calculated as the overall R^2 multiplied by the respective Shapley contribution of each variable group. The decomposition again highlights the importance of traditional variables.

When looking at the rolling window regression that goes back to the 1970s and 1980s, the R^2 remains stable and low at about 3%. The overall R^2 then exhibits a monotonic increase over the entire period, reaching a maximum value of approximately 30% in the final sample.

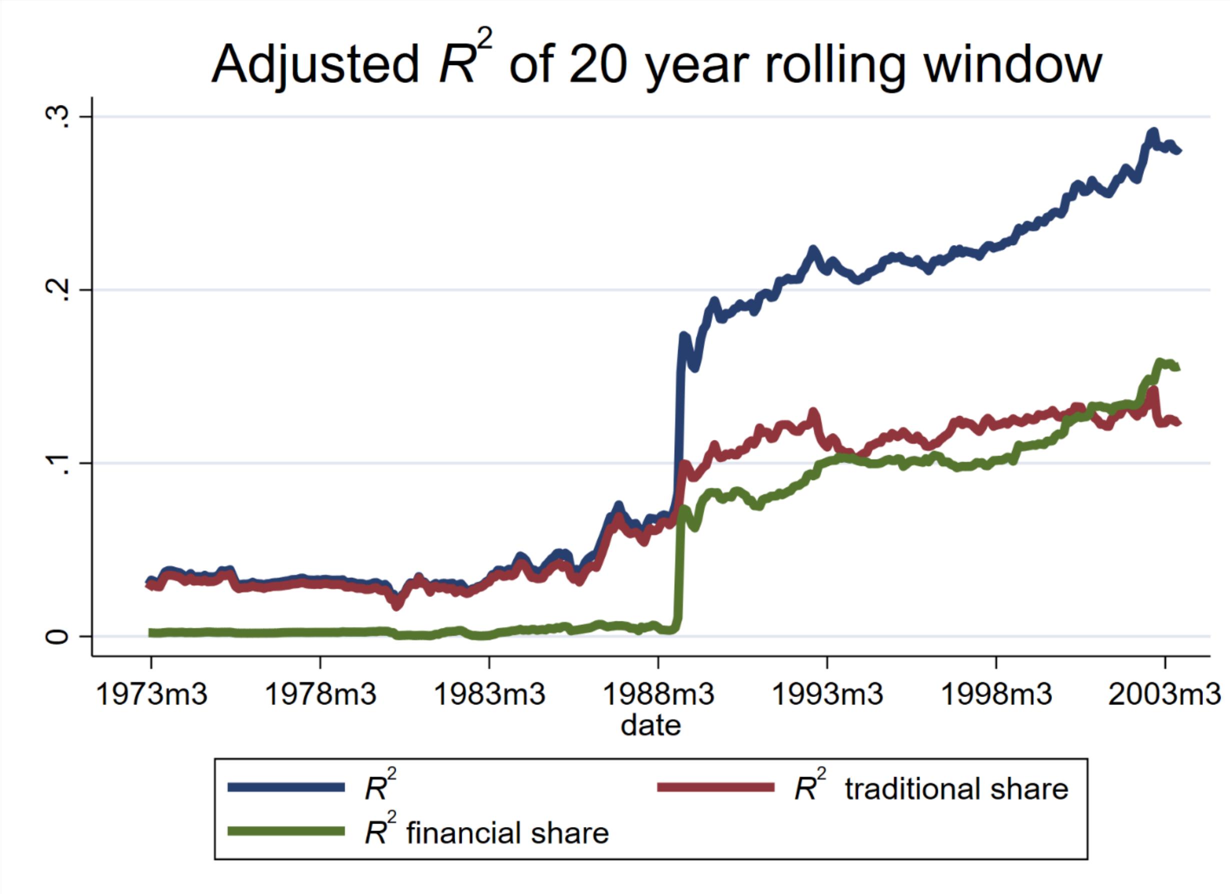
The overall R^2 exhibits a discontinuous jump around 1988, which is when the global financial crisis period begins to be included in the 20-year rolling regression sample. We can understand this jump better by looking at the decomposition of the R^2 into traditional and financial variables. The explanatory power of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$) starts to improve when the sample start date is around 1983, which corresponds to the late 1990s and early 2000s entering the sample. The R^2 of the traditional and monetary variables exhibits a fairly stable upward trend, rising from about 5% around 1983 to approximately 15% at the end of the sample, without a dramatic change when the 2008 sample is included.

On the other hand, the risk variables contributed almost nothing to the overall R^2 before 2008. There is then a sharp increase in the risk variable contributed R^2, reaching about 8% when the 2008 sample is included. During the same period, the R^2 of the traditional variables is about 10%. Thereafter, financial variables continue to explain just under half of exchange rate variation, eventually slightly overtaking traditional variables by the end of the sample. The jump in the R^2 in 2008 comes primarily from the contribution of the financial variables, but that jump is not the main story in the overall improvement of model fit.

We highlight the importance of the traditional variables inside and outside of the 2008-2009 Global Financial Crisis with two additional analyses.

[8] In Appendix B Table A4 and A5, we plot also the results for all individual currency pairs and also report the *F*-statistics for the null of no joint explanatory power of the right-hand-side variables.

Figure 3: R^2 of 20-year rolling window regressions of equation (1)



Note: The figure reports the R^2 in equation (1) with a 20-year rolling window regression. Shapley (R^2 traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$). Shapley ($R^2$ financial) reports the same Shapley value decomposition for $\Delta RISK_t$. 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.

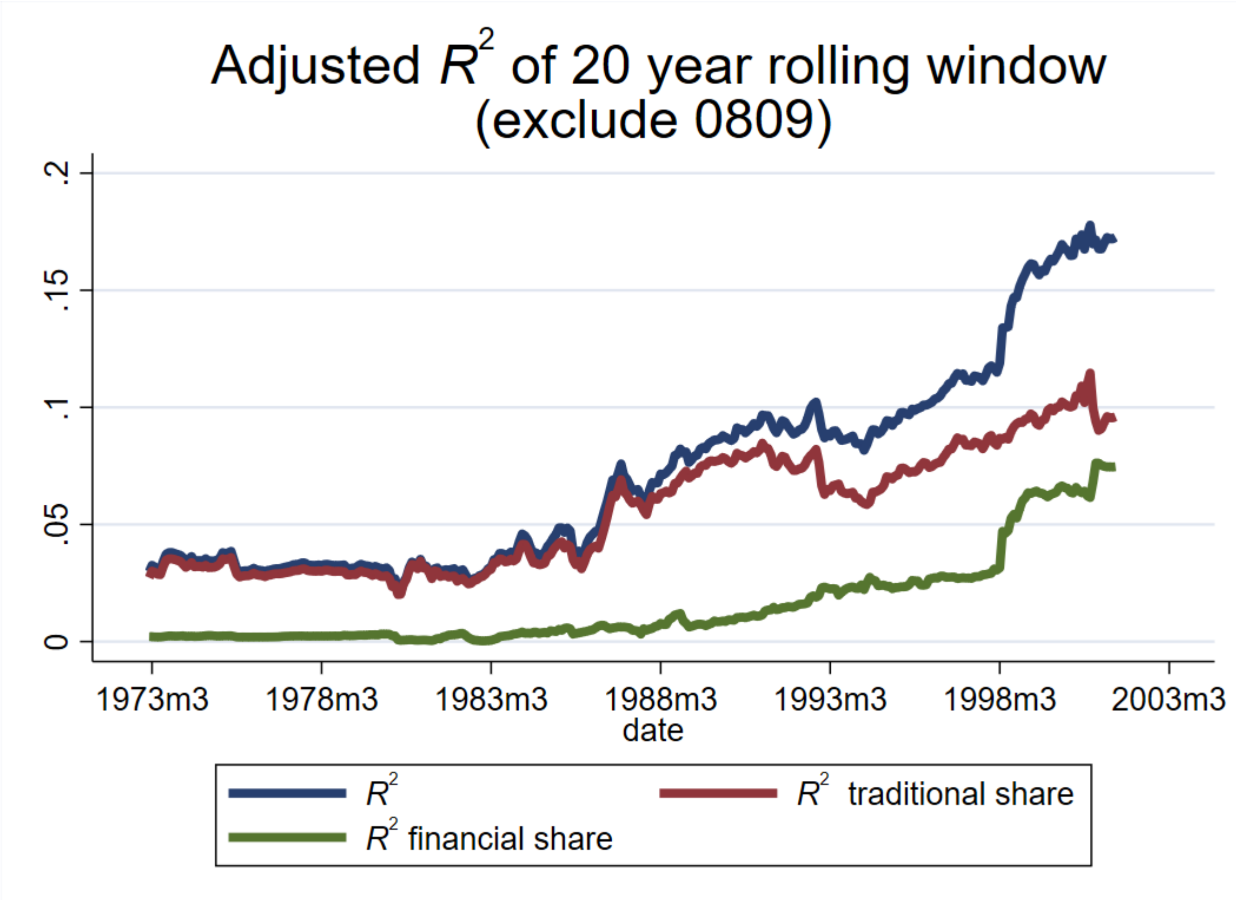
First, in Figure 4, we replicate the in-sample R^2 analysis from Figure 3 but exclude the 2008–2009 period (the country-by-country version is reported in Appendix Figure A9). This figure yields two key insights. First, removing the 2008–2009 crisis period reveals a much clearer picture of the monotonic improvement in explanatory power since the 1980s. Notably, there is no dramatic jump in R^2 around 1988 (which corresponds to the 20-year sample from 1988 to 2010, omitting 2008–2009). This steady improvement is a central finding of our paper. Second, examining the Shapley R^2 decomposition, we find that the traditional variables account for a high share of the R^2 and consistently exceeding the contribution of the financial variables throughout the entire sample. This suggests that the 2008–2009 data inflates the apparent importance of financial variables; as one might intuitively expect, the financial market dynamics during this crisis represent an outlier. Once this period is removed, the relative contribution of the financial variables notably diminishes. By re-estimating the Table 1 regression without the 2008–2009 data, we show that the contribution of traditional variables increases by 7 percentage points.[9]

Second, in Table 4, we report the baseline panel fixed-effects specification from equation (1) restricted exclusively to the 2008–2009 sample. We focus on the panel fixed effect version, given that the window contains only twenty-four monthly observations per currency pair. Consistent with Table 1, all point estimates maintain their expected signs (though the coefficients on foreign inflation and the convenience yield are estimated with less precision). The overall R^2 is high at 65%. Most importantly, as shown in the second to last row, the Shapley contribution of traditional variables stands at 54%, exceeding the 42% contribution attributable to financial variables. Taken together, these two exercises demonstrate that in the post-2000 period, traditional macroeconomic factors remain crucial for understanding exchange rate dynamics during both tranquil and crisis regimes.

The analysis shows that while exchange rate models performed poorly in earlier decades, their explanatory power has steadily improved over time. While recent literature—such as Lilley et al. (2022)—emphasizes the role of risk and financial variables in the post-2008 period, our findings indicate that traditional macroeconomic variables are a primary driver of exchange rate fluctuations. We next confirm this finding using an out-of-sample analysis.

[9] See Appendix C for a full analysis of the sample that excludes 2008-2009.

Figure 4: R^2 of 20-year rolling window regressions of equation (1), excluding 2008-2009



Note: The figure reports the R^2 in equation (1) with a 20-year rolling window regression, excluding 2008 and 2009. Shapley (R^2 traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables (Δr_t, Δr_t^*, π_t,π_t^*, $\frac{TB}{GDP_t}$). Shapley (R^2 financial) reports the same Shapley value decomposition for $\Delta RISK_t$. 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.

Table 4: Baseline regression, but restricted to 2008-2009

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

	Panel (FE)
Δr_t	-2.70***
	(-4.55)
Δr_t^*	2.23***
	(10.24)
π_t	-1.11**
	(-2.19)
π_t^*	0.11
	(0.66)
$\frac{TB}{GDP_t}$	-3.24**
	(-2.18)
$\Delta RISK_t$	-0.02***
	(-10.60)
$\Delta \eta_t$	-0.92
	(-1.62)
q_{t-1}	-0.11***
	(-3.60)
N	168
F	18.40
Adjusted R^2	0.65 (within)
Shapley (Traditional)	54%
Shapley (Financial)	42%

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 2008 to Dec 2009. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. Standard errors are Newey West standard errors with 12 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 12 lags. 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). Shapley (traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables (Δr_t, Δr_t^*, π_t,π_t^*, $\frac{TB}{GDP_t}$). Shapley (financial) reports the same Shapley value decomposition for $\Delta RISK_t$ and $\Delta \eta_t$.

c. Meese Rogoff out-of-sample fit from 1973

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

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

where each of the α and β are estimated using sample from period $t-1$ to period $t-240$. In each 240-month 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.[10]

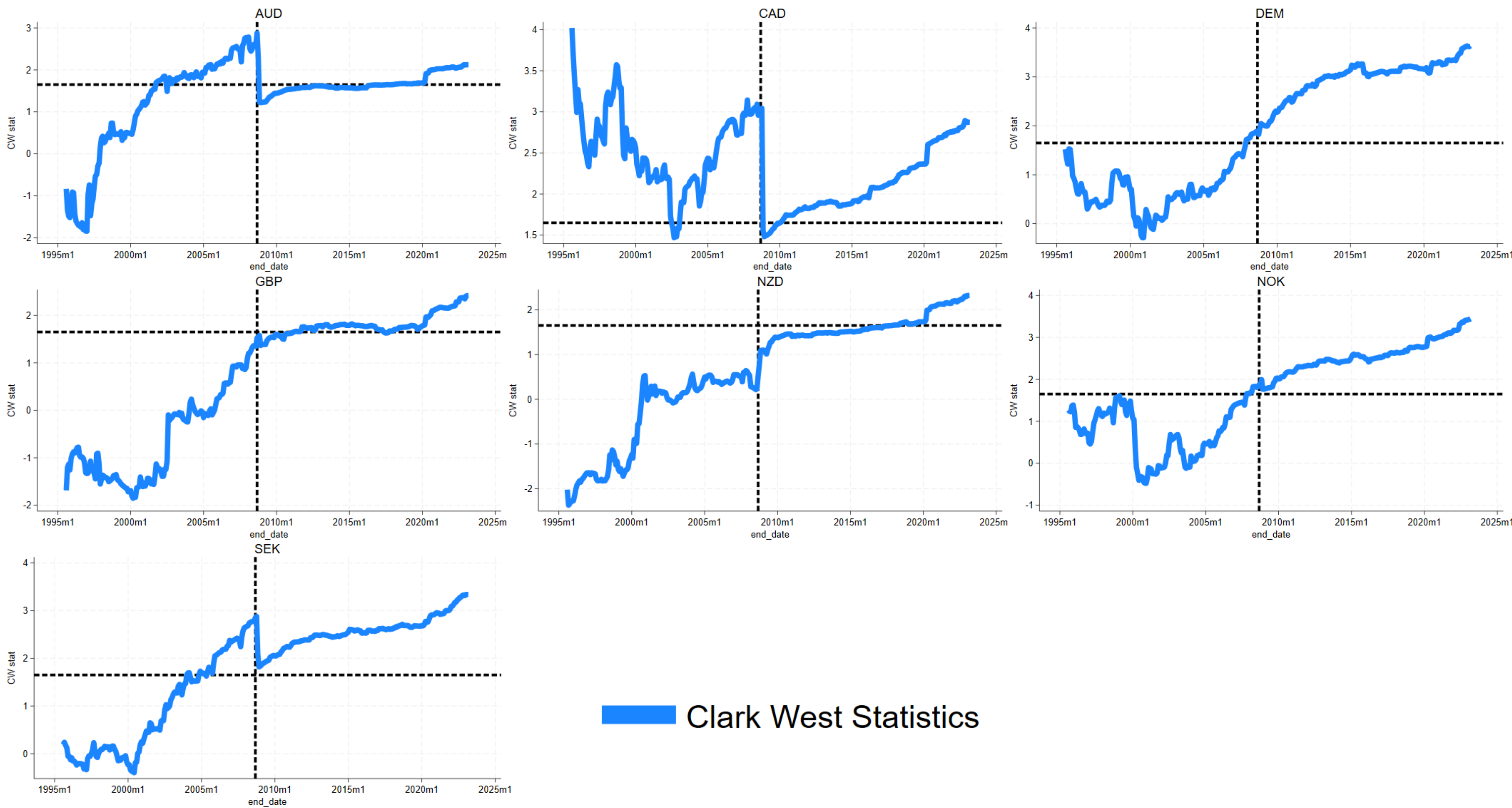
Since our sample starts from March 1973, we use the sample from March 1973 to February 1993 to estimate the first regression and generate the first prediction in March 1993. The sample is kept fixed for 240 months so regression coefficients of the second prediction in April 1993 are estimated using data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

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

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. 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. The Clark West statistics are computed using Newey West standard errors with 12 lags.

[10] While the Meese-Rogoff exercise is often referred to as a "forecasting" test, it is in fact not that, but instead a test of out-of-sample fit. Meese and Rogoff's seminal paper treated the nominal exchange rate as a stationary variable, and the random walk was not nested in the economic model. Out-of-sample fit should properly be considered a test for overfitting or data mining – of which perhaps the early empirical models of the 1970s were guilty. It is less surprising that we find little difference between the in-sample and out-of-sample results given the longer time series over which we can fit the models. A longer estimation period makes it much less likely that overfitting accounts for the findings because parameter estimates are more stable.

Figure 5: 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. The horizontal dash line indicates the *t*-statistics value of 1.65. The vertical dash line indicates the date of Sept 2008. Standard errors are Newey West standard errors with 12 lags.

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 random walk model produces a smaller mean prediction error than the exchange rate model, a well-documented feature of the "exchange rate disconnect".

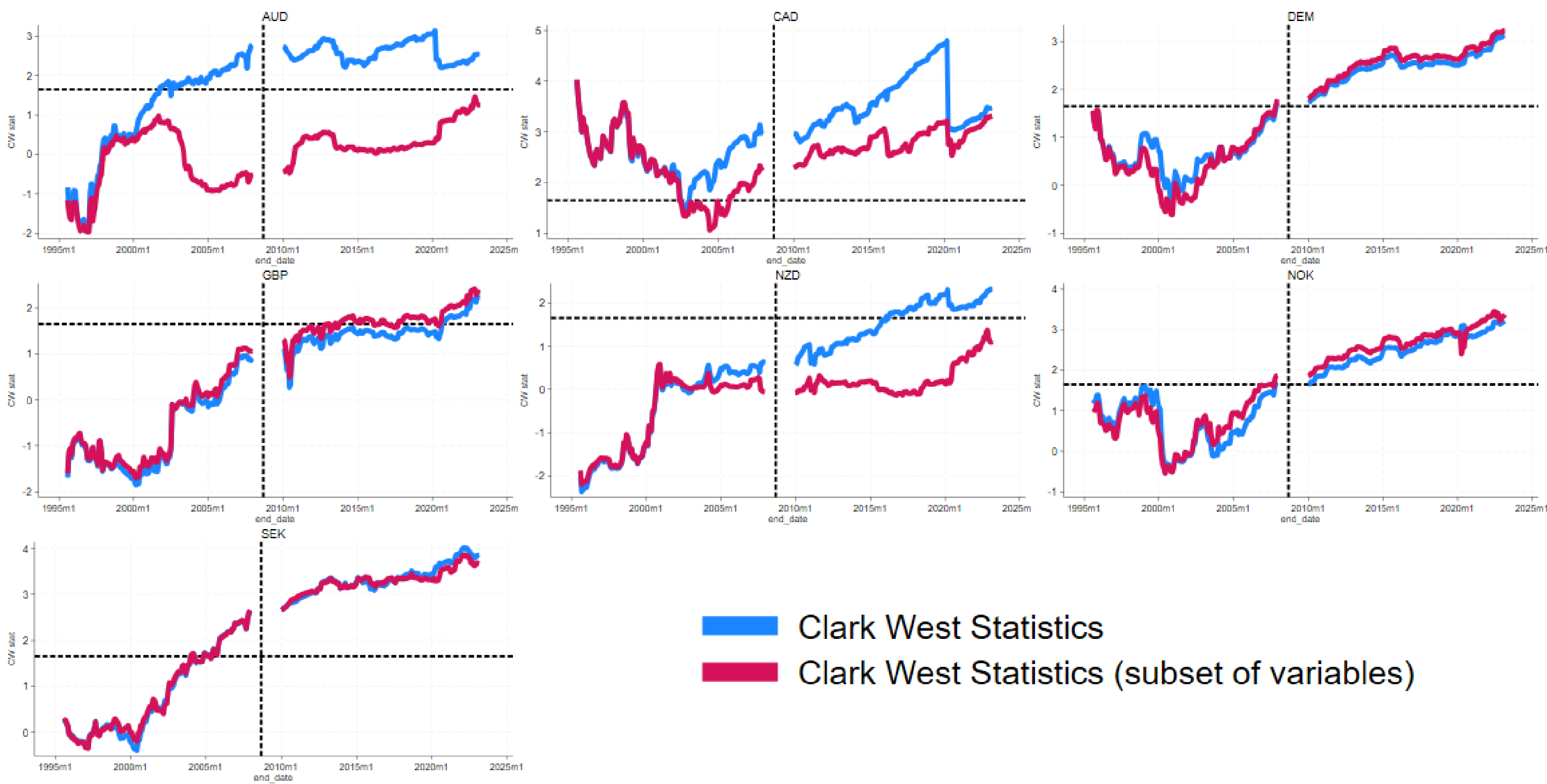
However, as time goes on, we find that the performance of the exchange rate model gradually increases. For all currencies, the exchange rate model beats a random walk model by the end of the sample significantly at the 1% level for one-sided test (Clark-West statistic above 1.96).

Although the exact point at which the exchange rate model significantly outperforms the random walk varies by country, there is strong evidence that the model's superiority to the random walk began prior to the Global Financial Crisis. The vertical dashed lines denote September 2008, marking the collapse of Lehman Brothers. For five currencies (AUD, CAD, DEM, NOK, and SEK), the exchange rate model provided a significantly better out-of-sample fit than the random walk prior to 2008—as early as 2003 for the AUD and SEK. The exceptions are GBP and NZD, for which the model only began to outperform the random walk around 2008 and 2009; however, both currencies still exhibited consistent performance improvements starting in 2000.

In Figure 6, we revisit the Meese-Rogoff out-of-sample fit exercise, excluding the 2008–2009 period from both the training and prediction samples. The blue lines in Figure 6 track those in Figure 5, but reflect this sample exclusion. Consistent with previous results, the Clark-West statistics indicate that while the exchange rate model initially underperforms a random walk, it begins to outperform it in the later portion of the sample across most currency pairs.[11] For the NZD and GBP—which in Figure 5 only outperformed the random walk around September 2008—the model requires several additional years of data to achieve statistical significance, with improvements becoming evident by 2015 and 2020, respectively. The model's relatively weak performance regarding the GBP is likely attributable to heightened political uncertainty following the Brexit referendum and its aftermath, dynamics that our theory-guided empirical model is not designed to capture.

[11] Clark West statistics are the same for Figure 5 and Figure 6 before 2008 as they are not affected by the sample exclusion.

Figure 6: 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. The horizontal dash line indicates the *t*-statistics value of 1.65. The vertical dash line indicates the date of Sept 2008.

To isolate the contribution of the traditional variables, the red lines in Figure 6 plot the Clark-West statistics of a restricted model against the random walk baseline. Specifically, we estimate the full regression containing both traditional and financial variables, but we zero out the financial variables when generating the out-of-sample forecasts. The red lines demonstrate that even when excluding the financial risk variables and omitting the 2008–2009 crisis period, the exchange rate model consistently outperforms the random walk.

Overall, we find that traditional macro factors contribute to the rise of the explanatory power of the exchange rate model significantly and it is not an artifact of the global financial crisis.

d. 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 for these two currencies. The fit is still relatively poor post-1999 and 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 are unlikely to be good measures of the monetary stance.[12] Sterilized intervention introduces another channel of influence on exchange rates that is not accounted for in the baseline model. And, while 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. 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

[12] Relatedly, due to persistent deflation, Japan and Switzerland have been operating at the zero lower bound for an extended period, which could lead to a self-fulfilling equilibrium, as we will argue in Section 4. Cole et al. (2023) document that Japan has the lowest central bank credibility among the advanced economies they studied.

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 predicted by theory. For example, the foreign interest rate effect is estimated to be negative for JPY, 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 rate 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 two 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 of the predictions from the portfolio balance model, because sterilized intervention may significantly alter the supply of bonds of each currency that are outstanding.

Taken together, the evidence in this section overturns the conventional view that exchange rates are disconnected from macroeconomic fundamentals. Since the late 1990s, a standard model built around real interest rates, inflation, the trade balance, and measures of global risk and liquidity fits U.S. dollar exchange rates with the theoretically correct signs, economically meaningful R^2, and reliable out-of-sample power against the random walk — a sharp reversal of the poor, often wrong-signed fit the same variables produced in the "old days" of the 1970s through the 1990s. This good fit is not an artifact of the 2008–2009 crisis: traditional macroeconomic variables account for roughly half or more of the model's explanatory power whether the crisis period is included, excluded, or examined on its own. What remains unexplained is why a regression that failed so badly before the mid-1990s succeeds so convincingly afterward. We turn to this question next.

Table 5: Baseline regression for Swiss franc and Japanese Yen from Jan 1999 – Aug 2023

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

	CHF	JPY
Δr_t	-1.42***	-0.36
	(-4.39)	(-1.19)
Δr_t^*	0.86*	-0.04
	(1.75)	(-0.10)
π_t	-0.27**	-0.29***
	(-2.10)	(-2.91)
π_t^*	0.35	-0.05
	(1.49)	(-0.30)
$\Delta RISK_t$	-0.00	0.01***
	(-0.62)	(3.77)
q_{t-1}	-0.01	-0.01
	(-0.85)	(-0.92)
$\frac{TB}{GDP_t}$	-0.18	-0.32*
	(-1.05)	(-1.71)
$\Delta \eta_t$	-2.40***	-1.48
	(-2.92)	(-1.38)
N	296	296
F	4.32	3.17
Adjusted R^2	0.08	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.

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

We find that the fit of a standard exchange-rate model for U.S. dollar exchange rates gradually improves since the mid-1990s, but the model did not fit well in the 1970s-80s. 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 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.[13] 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.

We first review in the context of a simple open-economy model why failure to target inflation sufficiently strongly would lead to poor inference in the empirical model. Then, in the next sub-section, we present some evidence that the closer relation of the exchange rate to its fundamental determinants coincides with the increasing credibility of monetary policy.

a. Model with insufficiently strong inflation targeting

We illustrate how the relationship between exchange rates and real interest rates and inflation may break down when the Taylor Principle is not satisfied in a simple two-country open-economy New Keynesian model. We use the "canonical" three-equation model from Engel's (2014) *Handbook* chapter. The simple canonical model discussed here is intended only to illustrate the causes of the underlying breakdown of the empirical exchange rate model. But it is not rich enough to capture meaningful endogenous relationship of variables, such as trade balance and risk premium. In Appendix E, we construct a full New Keynesian DSGE model with a global intermediary that takes limited positions in home and foreign bonds. We illustrate regression results from model-generated data in Table 6 below, comparing the model's ability to replicate our strong empirical results when the Taylor condition holds, and the poor fit when the monetary policy maker does not react sufficiently to changes in inflation.

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

The three equations are

(3) $E_t s_{t+1} - s_t - i_t^R = \rho_t$, (4) $\pi_t^R = \delta(q_t - \tilde{q}_t) + \beta E_t \pi_{t+1}^R$, (5) $i_t^R = \sigma\pi_t^R + \alpha i_{t-1}^R + \varepsilon_t^R$

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. We assume $\delta > 0$.

Equation (3) is the financial market equilibrium condition. 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 an exogenous random variable. This exogenous variable might be a risk or liquidity premium, that is endogenous in the more sophisticated set-up in Appendix E.

Equation (4) is a version of the home minus foreign Phillips curve and 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 $\bar{q}_t$ represents factors that drive the "equilibrium" real exchange rate in the long run, such as 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.

The stability condition for this model is $\sigma + \alpha > 1$. Algebraically, stability means that two eigenvalues are greater than one, and the other is less than one. When the stability condition is satisfied, exchange rates and inflation are forward looking and depend on expected paths of the exogenous driving shocks, but there is also a backward-looking component because of interest-rate smoothing.

When the stability condition is not satisfied, there is not a unique solution. There will be two roots less than one, and only one greater than one. There is one solution in which a particular linear combination of the endogenous variables is perfectly predictable, where the parameters in the linear combination are related to the eigenvector associated with the "unstable" root less than one. But there can also be sunspot solutions. A sunspot variable in our 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. In the sunspot solutions, no linear combination is perfectly predictable.

Appendix E simulates a 2-country model with sticky prices and asset trade through a global financial intermediary, in the spirit of Gabaix and Maggiori (2015), Itskhoki and Mukhin (2021), Kekre and Lenel (2024b) and Cormun and De Leo (2026). The model is calibrated to standard parameters and features monetary shocks, markup shocks and demand shocks for both countries and shocks to the UIP equation. When the stability condition is not met, the model is solved with the Bianchi and Nicolò (2021) method. [14]

When the Taylor condition is satisfied, the model can generate realistic t-statistics and R^2 as observed in the data. The simulated regressions are summarized in Table 6 column (1). It shows that when the Taylor principle is satisfied, the real interest rate, inflation, risk variable and trade balance are highly significant, and the R^2 is similar to our empirical findings.

In the simulations where Taylor condition is not met (columns 2 and 3), the R^2 values are low, as in the actual pre-1999 data. Inflation is significant but has the wrong sign for both the home and foreign country. Real interest rates also have the wrong sign but are largely insignificant. The variables representing risk or liquidity, $\Delta\omega_t$, and the trade balance, $\Delta b / Y_t$ are not statistically significant.

As noted above, we consider two cases. In the first case (column 2), there is no sunspot shock.[15] In the second (column 3), there is a sunspot shock with a variance similar to the size of other shocks. The performance of the model in these two cases is similar. We present them both as a way to emphasize that it is failure of inflation targeting that is driving the results, not noise introduced by sunspots.

Consider shocks that drive up current and future inflation in the model. When monetary policy is credible, the current and future real interest rate rises and the currency should appreciate now, and then gradually return to its long-run PPP value. When the monetary policy is not credible or when Taylor

[14] Bianchi and Nicolò solve models with sunspots by appending auxiliary exogenous equations—containing the sunspot shock and the endogenous forecast error—to the original system. By setting the root of these auxiliary equations to be explosive, they supply the missing unstable eigenvalues needed to satisfy standard Blanchard-Kahn conditions, allowing indeterminate models to be solved using off-the-shelf software like Dynare.

[15] However, the model without the sunspot does have a feature some might find undesirable – that a particular linear combination of the endogenous variables is perfectly predictable one period in advance – so we show the result carries over to a model with a sunspot shock where we (arbitrarily) set the volatility to be the same as the volatility of shocks to the monetary policy rule.

condition is not met, the nominal interest rate adjustment is too small. The market recognizes that higher inflation now will not lead to the sustained increase in real interest rates needed to bring inflation back down, so the currency depreciates rather than appreciates in response to the shock. Inflation is also associated with an opposite sign when the credibility changes but even the sign of shocks to risk or convenience may have the wrong sign because the overshooting/reversion to PPP mechanism is undermined. For example, if global risk increases, markets should be willing to accept a lower expected return on dollar assets. Under inflation targeting, this is accomplished by a current appreciation but with an expectation of a depreciation as adjustment occurs. When the Taylor condition is not satisfied, the lower expected return on dollar assets would be delivered by an immediate and persistent depreciation of the dollar.

Table 6: Regression with simulated data

	Taylor principle is satisfied	Taylor principle is not satisfied sunspot s.d. = 0	Taylor principle is not satisfied sunspot s.d. = 0.001
	(1)	(2)	(3)
$\Delta(r)_t$	-8.79***	0.59	0.23
$\Delta(r^*)_t$	8.84***	-0.57	-0.25
$(\pi)_t$	-2.31**	4.79***	4.40***
$(\pi^*)_t$	2.28**	-4.81***	-4.38***
$\Delta\omega_t$	-5.17***	-0.87	-0.80
$\Delta b_t/Y_t$	-2.76***	0.01	0.04
N	280 each, simulated 1000 times	280 each, simulated 1000 times	280 each, simulated 1000 times
Adjusted R^2	0.40	0.09	0.08

Note: median t-statistics are reported. * p<0.1, ** p<0.05, *** p<0.01. See Appendix E for the details of the model and simulation.

b. Evidence to support an inflation credibility channel

We provide three pieces of causal evidence for the inflation credibility channel. In the first exercise, we estimate country-by-country Taylor-rule equations and demonstrate that an improvement of credible inflation targeting rule is associated with structural breaks in the relationship between exchange rates and macro variables with a Difference in Difference (DiD) setup. Second is an instrumental variable (IV) exercise that shows an increase in a measure of central bank credibility is associated with a better fit of the exchange rate regression in equation (1). And finally, we provide a high frequency identification exercise to substantiate that a higher-than-expected inflation announcement is associated with an increase in expected future interest rate and dollar appreciation.

Before we dive into the causal evidence, it is helpful to visualize the slow-moving regime changes by looking at the *t*-statistics for the monetary variables in the upper 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.

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. As we have noted above, the other countries only adopted inflation targeting beginning in the 1990s. The coefficients on the real interest rates for those countries are generally estimated to be insignificant in the earlier part of the sample. However, 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. 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. 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 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., only in the latter half of the period does high inflation over the previous 12 months convince markets that the Fed would be tightening, leading to a stronger dollar.

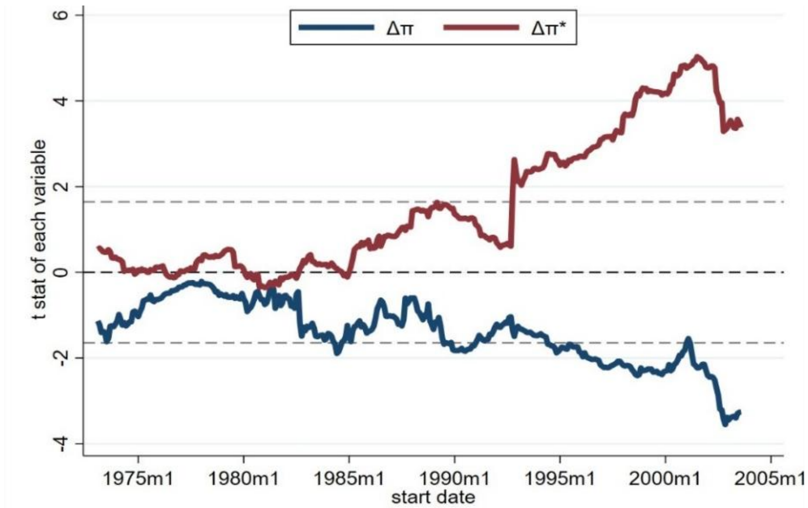
The findings for the measure of expected inflation in the foreign countries are similar, except that it is not until later windows (compared to the coefficient on U.S. inflation) that the estimates become significantly positive as the model would. This again is consistent with the observation that these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

In the baseline specification, we posited that the current level of inflation may measure the market's expectation of future monetary policy. When inflation is high, agents expect monetary policy to tighten, so the currency appreciates. We might interpret the change of inflation, $\pi_t - \pi_{t-1}$ as the "surprise" in inflation at time *t*. As Atkeson and Ohanian (2001) document, inflation in the previous 12 months is the best predictor of future inflation. In period $t-1$, the market's expectation of inflation for the next period is π_{t-1}, so the change in inflation proxies for the news this period about inflation. We estimate the baseline regression but use the change in inflation in the U.S. and the foreign country rather than the level of inflation. The full results are similar to the baseline regression and are reported in Appendix Table A6.

The estimated coefficients on U.S. and foreign inflation change measure the credibility of monetary policy. Taking into account the other determinants of the exchange rate, these coefficients indicate whether the surprise in inflation leads to an appreciation or depreciation of the currency. When U.S. inflation is unexpectedly high, the currency should appreciate if monetary policy is credible, so the coefficient should have a negative sign, and vice-versa for foreign inflation.

Figure 7 plots the *t*-statistics for the two variables in the 20-year rolling regressions. The horizontal dashed line is the critical value for 5% significance level in a one-sided test. By this metric, monetary policy was not credible in the 1980s but became increasingly credible in the 1990s and after 2000.

Figure 7: t-statistics on estimated coefficients of $\Delta\pi_t$ and $\Delta\pi_t^*$ in 20-year rolling regressions



Note: The figure reports t statistics for coefficients of $\Delta\pi_t$ and $\Delta\pi_t$. X-axis corresponds to the start date of the rolling window regression. The grey dash line represents the value of 1.65, the critical value for 5% significance level for one-sided test.

1: Estimated Switching Taylor Rules and Structural Breaks of Exchange Rate Regression

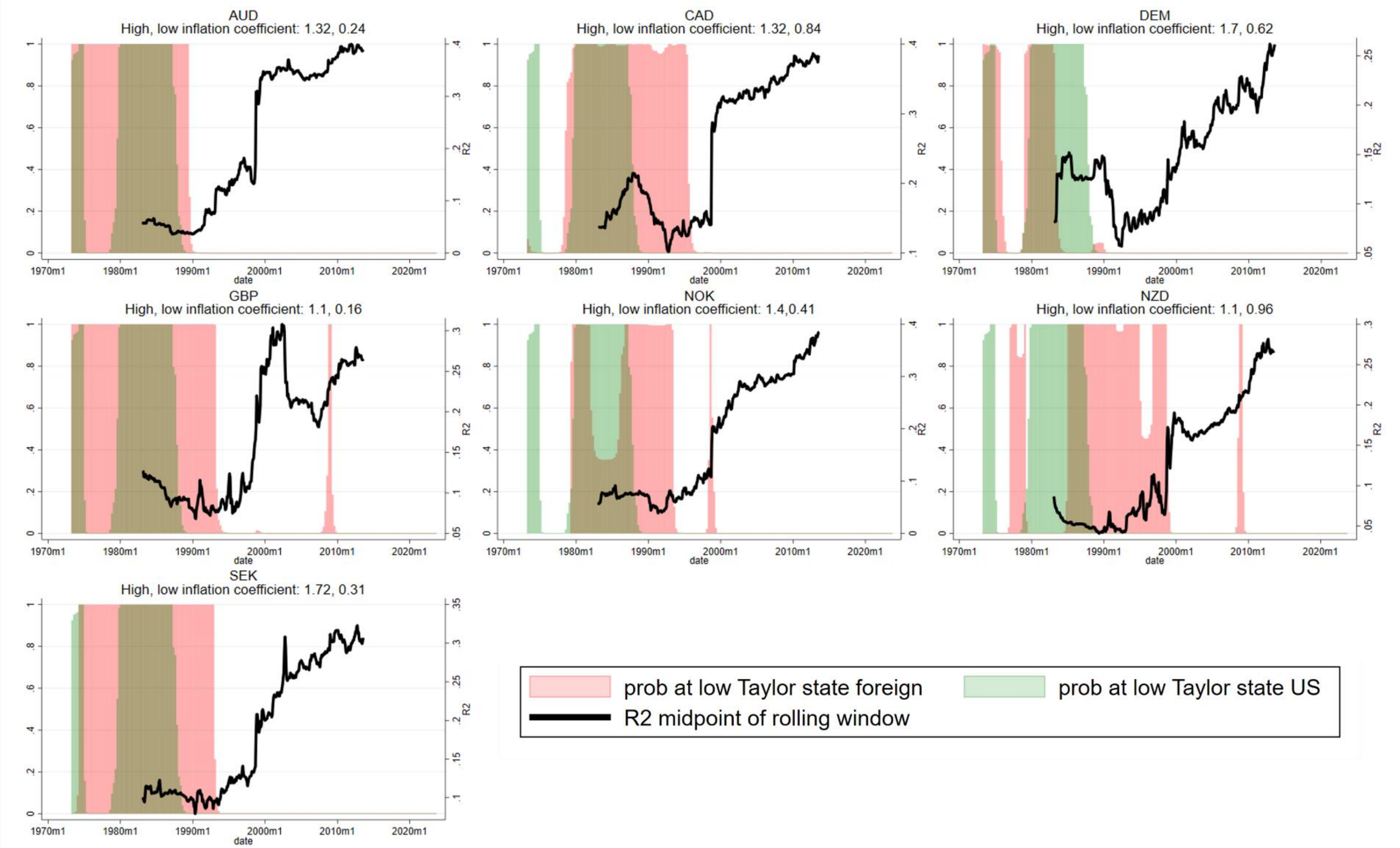
In the first exercise, we present new evidence of the evolving monetary policy regimes by estimating Taylor rules in the U.S. and other countries. We then provide evidence that a credible inflation policy changes the exchange rate regression fit.

Estimated Switching Taylor Rules. There is an extensive literature that investigates the increasing credibility of U.S. monetary policy in the late 20th century. Here, we use 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:

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

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.

Figure 8: 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.

We are interested in the coefficient $\phi_{\pi j}$ - credible policy implies the stability condition will be satisfied if $\phi_{\pi j} > 1$. We estimate a switching model for the monetary rule (6) 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 ($\phi_{\pi j} > 1$) to the low state to be zero.

The model is estimated by maximum likelihood. The smoothed probabilities of the states in which $\phi_{\pi j} < 1$ are reported in Figure 8. Supplementary Appendix D Figure A11 reports results when the stable state is not constrained to be an absorbing state. The general findings are not changed.

In Figure 8, the shaded green areas, which are the same in all panels, are the probability that the state is such that $\phi_{\pi j} < 1$ for the U.S.. The shaded red areas give the probability that the non-U.S. country *j* (of the subfigure title) 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. The estimated values of $\phi_{\pi j}$ in each regime are reported in the subtitles.

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.

Also plotted in Figure 8 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 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 monetary policy conduct has changed.

Structural Breaks of Exchange Rate Regression. To further strengthen this point for monetary credibility and exchange rate relationship, we conduct a simple event study analysis and a difference-in-differences analysis.

For each foreign country, we take the earlier date of either when the country implements inflation targeting (a de jure cutoff) or the estimated Taylor coefficient first becomes greater than one during the 1990s (a de facto cutoff) as an event date cutoff point that a country transits to a credible monetary regime. We estimate a two year panel fixed effect regression with only monetary variables for the two years before the cutoff and the two years after the cutoff to investigate how the exchange rate and interest rate relationship evolves around the regime change. Since we are exploiting the variations in foreign countries' inflation policies, we *designate each foreign currency as the base currency* against all other currencies, including the U.S. dollar and estimate the following regression:

(7) $\Delta s_{i,t} = \alpha + \beta_1 \Delta r_{i,t} + \beta_2 \Delta r^*_{i,t} + \beta_3 \pi_{i,t} + \beta_4 \pi^*_{i,t} + \beta_5 q_{i,t-1} + u_{i,t}$

The regression results are reported in Table 7. The variation that we exploit here is how the interest rate of the country affects its currency value relative to a panel of other currencies before and after the event date. It is clear from Table 7 that the transition to a credible monetary regime matters for the exchange rate regression. The coefficients on the home interest rate, β_1 in equation (7) are estimated to be insignificant for all country panels before the cutoff date. Two out of the seven coefficients are estimated to be positive, the opposite sign of what a monetary model would predict. On the other hand, after the cutoff date, all the coefficients on home interest rate become significantly negative at the 5 percent level, which aligns well with monetary model's prediction.[16]

Table 7: Two-year regression of exchange rate and interest rate before and after inflation targeting.

Regression: $\Delta s_{i,t} = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r^*_{i,t} + \beta_3 \pi_t + \beta_4 \pi^*_{i,t} + \beta_5 q_{i,t-1} + u_{i,t}$

Base currency	Panel for AUD	Panel for CAD	Panel for DEM	Panel for GBP	Panel for NZD	Panel for NOK	Panel for SEK
β_1, two-year sample before the cutoff	-1.04 (0.79)	-0.62 (0.86)	0.08 (0.53)	-0.49 (0.71)	0.30 (0.72)	-0.30 (0.25)	-0.36 (0.22)
β_1, two-year sample after the cutoff	-1.35*** (0.33)	-0.61** (0.24)	-1.47** (0.56)	-2.98*** (1.03)	-0.71** (0.26)	-0.59*** (0.21)	-0.46** (0.22)
Cutoff date	1989 Dec	1991 Jan	1990 Jan	1992 Oct	1989 Dec	1992 Mar	1993 Jan

Note: The table reports the coefficients on home interest rate for each column head currency as the home currency. Driscoll Kraay (1998) standard errors with 4 lags in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Observation for each panel regression is 168.

We further exploit the cutoff dates presented in Table 7 as treatment dates for the adoption of credible inflation-targeting (IT) policies. These dates serve as a natural benchmark for identifying structural breaks in the relationship between exchange rates and macroeconomic fundamentals in our baseline regressions for the U.S. dollar exchange rate. Specifically, we perform a staggered difference-in-differences analysis and structural break tests to evaluate the exchange rate's sensitivity to

[16] In Appendix Table A11, we also estimate the model for four to two years prior, and six to four years prior to the cutoff as a placebo test. We find that the coefficients on the home interest rate remain insignificant or significant with opposite signs.

macroeconomic variables changes following the implementation of a credible inflation policy. Countries that adopt inflation targeting are classified as the treatment group, while those that have not adopted the policy by time *t* serve as the control group.

To formalize this analysis, we construct a country-specific inflation treatment dummy variable, $IT_{i,t}$, which equals one if the observation occurs after the treatment cutoff date in Table 7 for country *i*, and zero otherwise. To compare with the potential structural change induced by the Global Financial Crisis, we also introduce a dummy for post 2008 sample, $I_{post08,t}$. We estimate the following panel regression model over the full sample period from March 1973 to August 2023:

$$(8)\Delta s_{i,t} = \alpha_i + \beta_1 \Delta r_t + \beta_2 \Delta r^*_{i,t} + \beta_3 \pi_t + \beta_4 \pi^*_{i,t} + \beta_5 \frac{TB}{GDP_t} + \beta_6 \Delta RISK_t + \beta_7 q_{t-1} + \beta_8 I_{post08,t}$$

$$+\beta_9 IT_{i,t} + IT_{i,t} \times [\beta_{10} \Delta r_t + \beta_{11} \Delta r^*_{i,t} + \beta_{12} \pi_t + \beta_{13} \pi^*_{i,t} + \beta_{14} \frac{TB}{GDP_t} + \beta_{15} \Delta RISK_t + \beta_{16} q_{t-1}]$$

$$+I_{post08,t} \times [\beta_{17} \Delta r_t + \beta_{18} \Delta r^*_{i,t} + \beta_{19} \pi_t + \beta_{20} \pi^*_{i,t} + \beta_{21} \frac{TB}{GDP_t} + \beta_{22} \Delta RISK_t + \beta_{23} q_{t-1}] + u_{i,t}$$

This specification builds on our baseline exchange rate regression in equation (1) by interacting all macroeconomic variables with the treatment dummy and 2008 dummy, allowing us to capture any structural shift in their effects post-policy adoption.

The regression results, reported in Table 8, provide strong evidence of a structural break associated with the implementation of credible inflation-targeting policies. In the pre-treatment period, only the U.S. interest rate is statistically significant, with a coefficient of -0.70, suggesting that exchange rate movements respond primarily to domestic interest rate. Other macro variables, including foreign interest rates, inflation, global risk, and trade balances, do not show significant effects prior to the treatment.

In contrast, the interaction terms—representing the effects of macro variables after the adoption of IT—are largely and significantly different from zero. For example, the interaction between the inflation treatment dummy and domestic inflation yields a much larger coefficient of -0.23, significant at the 1% level, indicating a stronger exchange rate response post-treatment. Similarly, the foreign inflation term becomes highly significant only after treatment (coefficient = 0.21, t = 3.56), as do both domestic and foreign interest rate variables. Global risk also shows a significant post-treatment effect, with a negatively significant coefficient, while the treatment dummy interacted with the trade balance ratio is statistically significant at the 5% level.

Table 8: Regressions with interaction with treatment dummy variables.

Panel fixed effect regression					
Δr_t	-0.70***	$IT_{i,t} \times \Delta r_t$	-1.51***	$I_{post08,t}\, \Delta r_t$	0.10
	(-4.33)		(-6.34)		(0.29)
Δr_t^*	0.01	$IT_{i,t} \times \Delta r_t^*$	0.54***	$I_{post08,t}\, \Delta r_t^*$	0.63**
	(0.06)		(3.21)		(2.30)
π_t	-0.03	$IT_{i,t} \times \pi_t$	-0.23***	$I_{post08,t}\, \pi_t$	-0.12
	(-0.80)		(-2.89)		(-1.35)
π_t^*	0.00	$IT_{i,t} \times \pi_t^*$	0.21***	$I_{post08,t}\, \pi_t^*$	-0.00
	(0.05)		(3.56)		(-0.06)
TB/GDP_t	-0.11	$IT_{i,t} \times TB/GDP_t$	-0.17**	$I_{post08,t}\, TB/GDP_t$	-0.31
	(-1.12)		(-2.31)		(-1.26)
$\Delta RISK_t$	-0.00	$IT_{i,t} \times \Delta RISK_t$	-0.02***	$I_{post08,t}\, \Delta RISK_t$	-0.00
	(-0.36)		(-3.32)		(-0.57)
q_{t-1}	-0.01	$IT_{i,t} \times q_{t-1}$	-0.01	$I_{post08,t}\, q_{t-1}$	0.00
	(-1.35)		(-1.25)		(0.97)
		$IT_{i,t}$	-0.00	$I_{post08,t}$	-0.01
			(-0.82)		(-1.37)
N	5337	Joint F test statistics of all the $IT_{i,t}$ interacted terms are zero (Chow test statistics)		10.25***	

Note: Driscoll Kraay (1998) standard errors with 12 lags in parentheses: * p<0.1, ** p<0.05, *** p<0.01. $I_{post08,t} = 1$ if year$\geq$ 2008. $IT_{i,t} = 1$ if time is after the cutoff date in Table 7.

Crucially, the F-test for the null hypothesis that all $IT_{i,t}$ interaction terms are jointly zero (the Chow test) is strongly rejected at the 1% significance level (F-statistic = 10.25). This rejection further confirms the presence of a structural break in the exchange rate determination process surrounding these policy implementation dates.

Conversely, the post-2008 interaction dummies are largely statistically insignificant across most macroeconomic variables, with the exception of the foreign interest rate, which is significantly positive. This suggests that, after controlling for the parameter shifts associated with the mid-1990s adoption of inflation targeting, there is no distinct structural change in the slope coefficients during the 2008 crisis.[17]

It is worth thinking about why the overall regression fit (R^2) improves substantially during 2008, as shown in Figure 3, if there is no structural change in the coefficients. During the 2008 crisis, there

[17] In Appendix Table A12, we provide the regression results without interacting 2008 to focus on the shift coming from inflation targeting. We also provide statistics in Table A13 to show that the structural break inflation targeting dates are more statistically plausible than a break in 2008.

was extreme variance in both the exchange rate and the explanatory macro variables. For instance, the U.S. dollar appreciated by 8% relative to the euro in October 2008, marking the largest single-month appreciation for the dollar-euro exchange rate in our sample. Additionally, there were five instances of approximately a one-standard-deviation drop (50 basis points) in the monthly change of the U.S. 3-month interest rate during 2008. The standard deviation of the global risk variables during the 2008–2009 period was double that of non-crisis periods. The dramatic increase in R^2 is driven by the high variance in both independent and dependent variables; the extreme data points fall along the existing slope, amplifying the signal of the underlying relationship rather than indicating a structural change in the coefficient.

2: Instrumental variable regression

In the second exercise, we directly use a measure of central bank credibility to investigate how increases in credibility are associated with improvements in the performance of exchange rate regressions. We measure central bank credibility using the indices constructed by Romelli (2024), which are well-suited to our analysis for two key reasons. First, the index offers broad coverage, spanning a wide set of countries from 1973 onward, which aligns well with our sample period. Second, Romelli constructs the index for each central bank based on a comprehensive set of institutional features, including the independence of the governing board, autonomy in setting monetary and financial policy, the statutory goals of the central bank, restrictions on lending to the government, financial independence, and reporting and disclosure requirements. These dimensions provide a rich and consistent measure of central bank performance, independence, and credibility over time.

We then conduct a panel fixed effects regression of the change in the R-squared from rolling exchange rate regressions on the change in the central bank credibility index (CBI) over time:[18]

$$\Delta R^2_{i,t} = \alpha_i + \beta \Delta CBI_{i,t} + u_{i,t}$$

Since central bank credibility may respond endogenously to a country's own macroeconomic conditions, we employ an IV approach to address this concern. We adopt a peer-pressure IV strategy used by Acemoglu et al. (2019) and Jung et al. (2025). Specifically, we use the one-period lag of the simple average of CBI in the same region, excluding the country itself, as an instrument.

[18] The central bank index is available at annual frequency, we take the simple average of monthly R-squared to aggregate it to annual level. We account for the potential serial correlation of the errors using Driscoll Kraay (1998) standard error with 20 lags. We also report 3 lags in the Appendix Table A14.

The intuition behind the instrument is that countries may be incentivized—or feel pressured—to implement similar reforms when neighboring or economically linked countries successfully adopt those reforms. A relevant example in our context is the adoption of inflation targeting by New Zealand in 1990, which influenced Australia to adopt a similar framework in 1993. The identifying assumption is that the average central bank credibility in a region affects a country's own central bank reforms but does not directly affect the country's exchange rate regression fit, thus satisfying the exclusion restriction.

The first-stage regression is specified as follows:

$$\Delta CBI_{i,t} = \delta_i + \gamma \Delta \overline{CBI}_{-i,t-1} + e_{i,t}$$

where $\overline{CBI}_{-i,t-1}$ denotes the average lagged CBI in the region excluding country i itself. We follow the region classification of Acemoglu et al (2019).

The regression results are reported in Table 9. The first-stage F-statistic is 22, indicating that the instrument has sufficient strength to explain variation in the CBI. The second-stage coefficient on the instrumented change in CBI is estimated to be 2.03 and is highly statistically significant. From 1990 to 2023, the average CBI increased by 0.15, implying that the predicted R-squared would rise by approximately 30.5%. This estimate aligns closely with the improvement observed between Table 1 and Table 3. As a robustness check, in Appendix Table A14, we also report the case in which we use the lagged change in the central bank index for each country within the region as separate instruments and conduct an overidentification test. The IV regression yields similar results, and the null hypothesis that the instruments are jointly valid cannot be rejected.

Table 9: Instrumental variable regression

	Panel Fixed effect regression
$\Delta\widehat{CBI}_{i,t}$	2.03***
	(4.86)
First stage F statistic	22
N	210

Note: t-statistics in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Sample period is from 1973 to 2023. The left-hand side variable in the regression is the annual change of R-squared from the exchange rate regression. The right-hand side variable is the change of central bank index from Romelli (2025). The instrumental variable is the lagged value of the average central bank index in the region, excluding the country itself. Driscoll Kraay (1998) standard error with 20 lags is reported.

3: Response of Exchange Rates to Inflation Announcements

One clear indication of whether markets believe that central banks will react to higher inflation comes from the movements of exchange rates and interest rates in response to news about inflation. We make use of the event when national statistical agencies announce the measure of inflation for an earlier month and investigate the exchange rate and interest rate response on those dates and find support for our mechanism. 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 to increase short term 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.[19]

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. The inflation surprises are measured as:

$$inflation\ surprise_{i,t} \equiv (inflation\ announcement\ value - inflation\ survey\ value)_{i,t}$$

For our purposes, however, the survey data is not ideal as 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.[20]

In Table 10, we conduct four separate regressions to provide evidence for the inflation credibility channel. The sample period is January 1997 to April 2024.[21] First, we regress daily exchange rate change

[19] A closely related exercise is Clarida and Waldman (2008). Related predecessors are Engel and Frankel (1984) and Andersen, et al. (2003).

[20] 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 Andersen, 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, in contrast to what we find in this section, these papers do not find a significant relationship between foreign inflation surprise and exchange rate adjustment 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 7.

[21] The exact start date differs by currency. Appendix Table A2 reports sample period by currency. We adjusted for the time zone difference if the inflation announcement is made after New York noon time. As discussed above, we exclude the sample from 2008-2009 to avoid the results from being driven by the Global Financial Crisis.

on the inflation surprise measures of the U.S. and foreign countries on the inflation announcement dates (column (1)). Next, we use three-month overnight index swap rates (OIS rates) to be our proxy of future interest rate movements. We regress the one-day OIS rate change on the corresponding inflation surprise measures, one for the U.S. (column (2)) and one panel fixed effect regression for the foreign countries (column (3)). Finally, to establish the exchange rate-interest rate relationship on the inflation announcements dates, we run a two stage least squares regression of exchange rate change on OIS rate change and instrument the OIS rates change with both countries' inflation surprises, controlling for daily change of global risk measure and convenience yield to ensure that the error term is uncorrelated with the instruments.

Formally, the regressions are:

$$(9)\quad \Delta s_{i,t} = \alpha + \beta_1 (US\ inflation\ surprise)_t + \beta_2 (Foreign\ inflation\ surprise)_{i,t} + u_{i,t}$$

$$\Delta OIS_{US,t} = \alpha + \beta_1 (US\ inflation\ surprise)_t + u_{i,t}$$

$$\Delta OIS_{i,t} = \alpha + \beta_1 (Foreign\ inflation\ surprise)_{i,t} + u_{i,t}$$

And

$$\Delta s_{i,t} = \alpha_i + \beta_1\, \Delta\widehat{OIS_{US,t}} + \beta_2\, \Delta\widehat{OIS_{i,t}} + \beta_3 \Delta risk_t + \beta_4 \Delta\eta_{i,t} + u_{i,t}$$

where the OIS rates are instrumented by both of the inflation surprises.

Column 1 of Table 10 presents the regression results of one-day exchange rate changes on inflation surprises. The estimated coefficients support our hypothesis. When foreign inflation is announced to be higher than market expectations, the U.S. dollar depreciates relative to the foreign currency. This suggests that markets anticipate the foreign central bank will credibly respond by tightening future real interest rates. A coefficient of 0.76 implies that the U.S. dollar depreciates by 0.76 percent in response to a one percentage point annualized inflation surprise abroad. Conversely, a positive U.S. inflation surprise tends to lead to an appreciation of the U.S. dollar on announcement days.

Columns 2 and 3 report regressions where the dependent variables are the one-day changes in U.S. and foreign OIS rates, respectively. The results show that positive inflation surprises are significantly associated with increases in expected future short-term interest rates. This indicates that markets expect central banks to respond to inflation surprises by raising policy rates—either the overnight rate itself or a closely correlated rate—reflecting credible monetary policy responses.

Column 4 uses inflation surprises of both countries as instruments for changes in OIS rates. The first-stage regression (reported in Appendix Table A15) has an F-statistic of 43.41, indicating a strong instrumental relevance. To address the exclusion restriction, we control for changes in risk and convenience yields to ensure that the error term is uncorrelated with the instruments, thereby ensuring that inflation surprises affect the exchange rate only through their impact on expected future short-term interest rates. In the second stage, the results show that an increase in the U.S. OIS rate driven by inflation surprises is associated with an appreciation of the U.S. dollar. In contrast, an increase in the foreign OIS rate—driven by inflation surprises—is associated with a depreciation of the U.S. dollar. Taken together, these findings provide strong evidence that markets perceive central banks as credibly responding to inflation shocks, leading to immediate adjustments in both interest rates and exchange rates.[22]

To summarize, the empirical evidence provided in this subsection suggests that the inflation credibility is relatively weak for a large set of advanced economies 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.

Table 10: Daily regressions on the inflation announcement dates

LHS variable:	$\Delta s_{i,t}$	$\Delta OIS_{US,t}$	$\Delta OIS_{i,t}$		$\Delta s_{i,t}$
	(1)	(2)	(3)		(4)
U.S.	-0.61**	4.88***		$\Delta \widehat{OIS_{US,t}}$	-0.22***
surprise	(-2.16)	(2.58)			(-2.62)
Foreign	0.76***		3.09***	$\Delta \widehat{OIS_{i,t}}$	0.28***
surprise	(9.80)		(6.99)		(5.30)
				$\Delta risk_t$	0.002
					(0.34)
				$\Delta \eta_{i,t}$	0.01
					(1.01)
R^2	0.03	0.04	0.04		0.01
N	3490	326	2651		2353
First stage F statistics					43.41

Note: *t*-statistics in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1997 to Apr 2024 but excludes 2008-2009. Driscoll Kraay (1998) standard errors are reported for the panel regressions.

[22] Bianchi and Melosi (2017) make the case that uncertainty about future policy during the zero-lower bound period led to the beginnings of a period of fiscal dominance in the U.S., but the results of Table 10 suggest that the Fed still retained inflation-targeting credibility.

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 due to the impossibility of perfectly measuring any of the variables in the model – inflation expectations, expectations of future monetary policy, global risk and liquidity, as well as the equilibrium real exchange rate. The 2025-2026 period, which goes beyond our baseline sample presents interesting challenges. This period has introduced new sources of risk, some of which are specific to the dollar itself (such as the "liberation day tariffs of April 2025 which are examined by Kekre and Lenel (2026), or the Iran war.) The rising long-term bond yields of mid-2026 may suggest a new era of fiscal dominance, and there have been concerns expressed about the independence of the Fed during the Trump administration. There are other variables that have yet to be discovered by theory, or that have already been discovered but not included here. Still, the models fit better than you might have thought.

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Supplementary Appendix

Appendix A: Data

Table A1: 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), FRED and OECD Long-term interest rates	3 month rates 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) 10 year rates OECD Long-term interest rates https://www.oecd.org/en/data/indicators/long-term-interest-rates.html
Risk variables	FRED. Federal Reserve Board For Gilchrist and Zakrajsek (2012)	FRED: AAAFF, BAAFF, AAA10Y, BAA10Y 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
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 (Only used for post 1999 regressions)	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: 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
Country by country GDP	OECD Main Economic Indicators	For output gap estimation

Table A2: Sample start date for the inflation announcement regression

Currency	Start date
AUD	28 Jan 1997
CAD	21 Feb 1997
EUR	19 Apr 2001
GBP	16 Jan 1997
NOK	10 Mar 1998
NZD	14 Jul 1997
SEK	12 Mar 1997

Appendix B: Robustness and additional evidence of section 3 on empirical exchange rate model

Table A3: Baseline Table 1 with 5 lags standard error adjustment

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

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel (FE)
Δr_t	-1.12***	-1.66***	-2.34***	-1.37***	-0.92***	-1.85***	-1.90***	-1.47***
	(-4.00)	(-6.43)	(-8.13)	(-5.07)	(-2.91)	(-5.70)	(-5.71)	(-6.34)
Δr_t^*	1.12***	1.16***	2.21***	1.80***	1.06**	0.23	0.79*	0.92***
	(3.61)	(4.63)	(6.18)	(3.58)	(2.41)	(0.99)	(1.92)	(4.13)
π_t	-0.25***	-0.21	-0.69***	-0.33**	-0.45***	-0.21**	-0.57***	-0.34***
	(-2.61)	(-1.55)	(-4.99)	(-2.47)	(-3.59)	(-1.98)	(-5.29)	(-4.86)
π_t^*	0.03	0.14	0.53***	0.14	0.24*	-0.19	0.25**	0.15**
	(0.25)	(0.93)	(4.02)	(1.26)	(1.72)	(-1.29)	(2.54)	(2.05)
$\frac{TB}{GDP_t}$	-0.48***	-0.45***	-0.63***	-0.73***	-0.37*	-0.66***	-0.80***	-0.54***
	(-2.71)	(-3.39)	(-3.84)	(-3.31)	(-1.80)	(-3.23)	(-3.84)	(-4.03)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01***	-0.01***	-0.02***	-0.02***	-0.02***	-0.02***
	(-7.57)	(-12.36)	(-2.62)	(-4.64)	(-6.14)	(-6.03)	(-4.52)	(-7.30)
$\Delta \eta_t$	-1.92**	-2.33***	-0.86	-1.52*	-1.56**	-1.20*	-0.69	-1.38**
	(-2.16)	(-3.47)	(-0.66)	(-1.93)	(-2.09)	(-1.80)	(-0.74)	(-2.44)
q_{t-1}	-0.01	-0.01	-0.02**	-0.03***	-0.01	-0.03***	-0.01	-0.01**
	(-1.33)	(-1.59)	(-2.39)	(-3.13)	(-1.52)	(-2.64)	(-1.48)	(-2.18)
N	296	296	295	296	296	296	296	2071
F	17.23	30.19	17.21	12.93	14.56	17.71	13.61	23.25
Adjusted R^2	0.36	0.36	0.25	0.22	0.22	0.30	0.24	0.25(within)

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. Standard errors are Newey West standard errors with 5 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 5 lags. 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 A4: Table 2 (regressions with 10-year interest rates and lagged variables) with 5 lags standard error adjustment

LHS: Δs_t	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel	Panel (quarterly)
$\Delta r_t^{10y,US}$	-8.92***	-5.27***	-6.40***	-1.48	-7.23***	-5.03***	-5.90***	-5.63***	-7.77***
	(-9.45)	(-4.59)	(-6.27)	(-1.30)	(-5.92)	(-5.11)	(-6.22)	(-9.24)	(-6.45)
$\Delta r_t^{10y,Foreign}$	8.74***	6.90***	6.46***	1.94	6.25***	5.71***	5.94***	5.91***	8.37***
	(8.56)	(4.98)	(5.12)	(1.16)	(4.48)	(5.18)	(4.37)	(6.62)	(6.43)
π_t^{US}	-8.49***	-4.33***	-4.86***	-0.76	-7.17***	-3.57***	-4.86***	-4.75***	-7.07***
	(-8.25)	(-3.68)	(-4.23)	(-0.65)	(-5.43)	(-3.37)	(-4.77)	(-7.27)	(-6.20)
$\pi_t^{Foreign}$	8.10***	6.31***	5.01***	1.20	6.13***	5.46***	5.95***	5.50***	7.67***
	(7.90)	(4.58)	(3.54)	(0.68)	(4.35)	(4.53)	(4.40)	(6.17)	(5.77)
$\frac{TB}{GDP_t}$	0.03	-0.33**	-0.56***	-0.42*	-0.25	-0.54**	-0.59***	-0.39***	-1.11***
	(0.16)	(-2.15)	(-2.77)	(-1.78)	(-1.15)	(-2.37)	(-2.73)	(-2.72)	(-3.43)
$\Delta Risk_t$	-0.04***	-0.02***	-0.01***	-0.01***	-0.03***	-0.02***	-0.02***	-0.02***	-0.03***
	(-12.76)	(-8.22)	(-2.87)	(-3.79)	(-7.32)	(-5.43)	(-5.01)	(-8.53)	(-8.45)
$\Delta\eta_t$	-1.37	-2.58***	-1.08	-2.06***	-2.22***	-1.47**	-1.11*	-1.53***	-1.97**
	(-1.63)	(-3.06)	(-0.95)	(-2.79)	(-3.07)	(-2.22)	(-1.66)	(-3.09)	(-2.15)
q_{t-1}	-0.06***	-0.00	-0.03***	-0.05***	-0.06***	-0.02**	-0.03***	-0.02***	-0.04***
Lagged variables below	(-4.72)	(-0.37)	(-2.85)	(-3.30)	(-3.90)	(-2.10)	(-2.96)	(-3.73)	(-2.75)
$r_{t-1}^{10y,US}$	-2.19***	-0.31	-0.50	-0.32	-1.83***	-0.77**	-0.66**	-0.58***	-2.18***
	(-4.12)	(-0.83)	(-1.52)	(-0.70)	(-3.12)	(-1.97)	(-2.22)	(-3.63)	(-5.39)
$r_{t-1}^{10y,Foreign}$	1.67***	0.38	0.33	0.57	1.16***	0.81***	0.68***	0.53***	1.81***
	(3.91)	(1.14)	(1.27)	(1.23)	(2.83)	(2.62)	(2.71)	(3.86)	(5.13)
π_{t-1}^{US}	6.19***	3.87***	3.84***	0.06	5.01***	2.49**	3.68***	3.87***	4.13***
	(5.52)	(3.23)	(3.39)	(0.05)	(3.59)	(2.37)	(3.48)	(5.92)	(3.44)
$\pi_{t-1}^{Foreign}$	-6.58***	-5.89***	-4.36***	-0.71	-4.97***	-4.73***	-5.16***	-4.94***	-5.81***
	(-5.72)	(-4.13)	(-3.13)	(-0.39)	(-3.42)	(-3.82)	(-3.72)	(-5.40)	(-4.37)
$Risk_{t-1}$	-0.00**	-0.00	-0.00	-0.00	-0.00**	-0.00	-0.00	-0.00	-0.00***
	(-2.55)	(-0.78)	(-0.02)	(-1.52)	(-2.19)	(-0.38)	(-1.37)	(-1.43)	(-2.70)
η_{t-1}	-0.35	-0.76	-1.74***	-1.86**	-0.91*	-0.95*	-1.30***	-0.89***	-0.42
	(-0.45)	(-1.12)	(-2.80)	(-2.44)	(-1.74)	(-1.74)	(-3.28)	(-2.85)	(-1.03)
N	296	296	295	296	296	296	296	2071	692
F	46.57	17.68	16.90	12.66	18.37	19.36	21.31	60	43
Adjusted R^2	0.48	0.40	0.32	0.24	0.31	0.37	0.36	0.32 (within)	0.55 (within)

Note: t-statistics in parentheses. * $p<0.1$, ** $p<0.05$, *** $p<0.01$. Sample period is from Jan 1999 to Aug 2023. Standard errors are Newey West standard errors with 5 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 5 lags for monthly regression and 4 lags for quarterly regression. Shapley (traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$ and its lags). Shapley (financial) reports the Shapley value decomposition for $\Delta RISK_t$ and $\Delta\eta_t$ and its lags.

Table A5: Table 2 quarterly regressions country by country

LHS: Δs_t	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel (quarterly)
$\Delta r_t^{10y,US}$	-11.96***	-7.88***	-7.09***	-3.57*	-9.99***	-6.07***	-8.10***	-7.51***
	(-7.35)	(-3.84)	(-3.75)	(-1.97)	(-4.02)	(-3.42)	(-5.63)	(-6.16)
$\Delta r_t^{10y,Foreign}$	10.55***	11.38***	8.60***	4.67**	8.36***	7.17***	8.50***	8.21***
	(7.29)	(4.62)	(4.14)	(2.55)	(3.62)	(3.81)	(5.85)	(6.32)
π_t^{US}	-11.29***	-6.15***	-5.62***	-3.14*	-10.06***	-5.14***	-8.39***	-6.84***
	(-6.91)	(-3.17)	(-2.81)	(-1.82)	(-3.92)	(-2.95)	(-5.40)	(-5.92)
$\pi_t^{Foreign}$	9.05***	9.32***	6.16***	3.38*	8.32***	7.27***	9.57***	7.55***
	(5.69)	(4.14)	(2.82)	(1.72)	(3.26)	(3.74)	(6.39)	(5.73)
$\frac{TB}{GDP_t}$	1.11*	-1.13***	-1.37***	-1.19*	-0.54	-1.52***	-1.51***	-1.10***
	(1.97)	(-3.23)	(-2.73)	(-1.84)	(-0.97)	(-2.95)	(-3.46)	(-3.40)
$\Delta Risk_t$	-0.04***	-0.03***	-0.01*	-0.02***	-0.04***	-0.02***	-0.03***	-0.03***
	(-11.23)	(-8.02)	(-1.87)	(-4.32)	(-6.52)	(-3.64)	(-5.31)	(-8.37)
$\Delta\eta_t$	1.38	-0.22	-6.74***	-2.62	-3.95**	-4.26**	-1.39	-2.86***
	(0.65)	(-0.11)	(-2.91)	(-1.36)	(-2.46)	(-2.36)	(-1.09)	(-3.18)
q_{t-1}	-0.18***	0.01	-0.07***	-0.14***	-0.16***	-0.04*	-0.04*	-0.04***
Lagged variables below	(-4.30)	(0.49)	(-2.95)	(-2.92)	(-4.17)	(-1.98)	(-1.85)	(-2.92)
$r_{t-1}^{10y,US}$	-7.91***	-1.74**	-2.19***	-1.59	-5.41***	-1.92*	-3.06***	-1.96***
	(-3.97)	(-2.08)	(-2.82)	(-1.22)	(-3.40)	(-1.72)	(-4.06)	(-5.19)
$r_{t-1}^{10y,Foreign}$	6.29***	1.75**	1.53**	2.06	3.54***	2.10**	2.71***	1.70***
	(3.94)	(2.38)	(2.40)	(1.57)	(3.10)	(2.50)	(4.24)	(5.11)
π_{t-1}^{US}	3.43*	3.63*	2.50	0.62	3.74	2.54	3.56**	4.15***
	(1.90)	(1.75)	(1.49)	(0.28)	(1.42)	(1.44)	(2.39)	(3.50)
$\pi_{t-1}^{Foreign}$	-3.35**	-7.17***	-4.24**	-1.57	-4.73*	-5.54**	-6.37***	-5.80***
	(-2.08)	(-3.24)	(-2.07)	(-0.68)	(-1.84)	(-2.63)	(-4.08)	(-4.38)
$Risk_{t-1}$	-0.02***	-0.00*	-0.00	-0.01*	-0.01**	-0.00	-0.01***	-0.00**
	(-3.46)	(-1.70)	(-0.56)	(-1.94)	(-2.51)	(-0.07)	(-2.78)	(-2.17)
η_{t-1}	1.34	1.22	-4.84**	-3.03*	-1.78	-3.23*	-2.24**	-1.50
	(0.68)	(0.63)	(-2.45)	(-1.77)	(-1.05)	(-1.84)	(-2.17)	(-1.65)
N	99	99	98	99	99	99	99	692
Adjusted R^2	0.72	0.62	0.54	0.45	0.56	0.56	0.62	0.55 (within)
Shapley (Traditional)	51%	59%	70%	53%	43%	60%	63%	53%
Shapley (Financial)	45%	40%	26%	41%	50%	38%	36%	45%

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Q1 1999 to Q2 2023. Standard errors are Newey West standard errors with 5 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 5 lags for monthly regression and 4 lags for quarterly regression. Shapley (traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables (Δr_t, Δr_t^*, π_t,π_t^*, $\frac{TB}{GDP_t}$ and its lags). Shapley (financial) reports the Shapley value decomposition for $\Delta RISK_t$ and $\Delta\eta_t$ and its lags.

Table A6 With liquidity variables, with inflation change and 3m Govt rate

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

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect
Δr_t	-4.11***	-3.91***	-3.47***	-2.43***	-4.35***	-2.61***	-3.37***	-2.89***
	(-5.56)	(-5.96)	(-5.00)	(-3.63)	(-5.03)	(-3.46)	(-4.37)	(-4.57)
Δr_t^*	4.38***	3.52***	2.85***	3.47***	4.18***	0.07	2.41***	1.36**
	(5.56)	(5.00)	(3.45)	(4.58)	(4.91)	(0.23)	(2.75)	(2.16)
$\Delta \pi_t$	-3.62***	-2.66***	-1.76**	-1.57**	-4.33***	-0.78	-2.01**	-1.81***
	(-4.46)	(-3.73)	(-2.24)	(-2.10)	(-4.56)	(-0.96)	(-2.39)	(-2.61)
$\Delta \pi_t^*$	3.45***	2.74***	0.62	2.01**	3.56***	-0.41	1.59*	0.58
	(4.25)	(3.81)	(0.64)	(2.27)	(3.99)	(-1.09)	(1.76)	(0.88)
$\Delta RISK_t$	-0.04***	-0.02***	-0.01***	-0.01***	-0.03***	-0.03***	-0.02***	-0.03***
	(-10.14)	(-9.57)	(-3.80)	(-4.71)	(-8.26)	(-7.33)	(-6.55)	(-9.39)
q_{t-1}	-0.00	-0.01	-0.02*	-0.01	-0.01	-0.01	-0.01	-0.01
	(-0.57)	(-1.54)	(-1.96)	(-1.27)	(-0.92)	(-1.42)	(-0.57)	(-1.54)
$\frac{TB}{GDP_t}$	-0.30*	-0.34***	-0.38**	-0.30*	-0.15	-0.45**	-0.33*	-0.32***
	(-1.91)	(-3.16)	(-2.49)	(-1.77)	(-0.87)	(-2.53)	(-1.84)	(-2.73)
$\Delta \eta_t$	-2.16**	-2.53***	-1.26	-1.70**	-1.98***	-1.31*	-0.91	-1.62***
	(-2.42)	(-3.25)	(-1.27)	(-1.97)	(-2.72)	(-1.87)	(-1.29)	(-2.66)
N	296	296	295	296	296	296	296	2071
F	25.76	23.93	10.35	11.12	13.44	14.77	10.73	24.19
R2	0.42	0.40	0.22	0.24	0.27	0.29	0.23	
R2_adjusted	0.40	0.38	0.20	0.22	0.25	0.27	0.21	0.24 (within)

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, *t*-statistics are based on 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.

Table A7 With Bianchi Bigio Engel liquidity ratio, with inflation level and 3m Govt rate

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

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

Table A8: Quarterly regression with 10-year interest rates and lagged variables

LHS: Δs_t	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel
$\Delta r_t^{10y,US}$	-11.96***	-7.88***	-7.09***	-3.57*	-9.99***	-6.07***	-8.97***	-7.77***
	(-7.35)	(-3.84)	(-3.75)	(-1.97)	(-4.02)	(-3.42)	(-6.44)	(-6.45)
$\Delta r_t^{10y,Foreign}$	10.55***	11.38***	8.60***	4.67**	8.36***	7.17***	9.12***	8.37***
	(7.29)	(4.62)	(4.14)	(2.55)	(3.62)	(3.81)	(5.51)	(6.43)
π_t^{US}	-11.29***	-6.15***	-5.62***	-3.14*	-10.06***	-5.14***	-9.26***	-7.07***
	(-6.91)	(-3.17)	(-2.81)	(-1.82)	(-3.92)	(-2.95)	(-6.00)	(-6.20)
$\pi_t^{Foreign}$	9.05***	9.32***	6.16***	3.38*	8.32***	7.27***	10.38***	7.67***
	(5.69)	(4.14)	(2.82)	(1.72)	(3.26)	(3.74)	(5.79)	(5.77)
$\Delta Risk_t$	-0.04***	-0.03***	-0.01*	-0.02***	-0.04***	-0.02***	-0.03***	-0.03***
	(-11.23)	(-8.02)	(-1.87)	(-4.32)	(-6.52)	(-3.64)	(-6.07)	(-8.45)
q_{t-1}	-0.18***	0.01	-0.07***	-0.14***	-0.16***	-0.04*	-0.02	-0.04***
	(-4.30)	(0.49)	(-2.95)	(-2.92)	(-4.17)	(-1.98)	(-0.85)	(-2.75)
$\frac{TB}{GDP_t}$	1.11*	-1.13***	-1.37***	-1.19*	-0.54	-1.52***	-1.41***	-1.11***
	(1.97)	(-3.23)	(-2.73)	(-1.84)	(-0.97)	(-2.95)	(-3.24)	(-3.43)
$\Delta\eta_t$	1.38	-0.22	-6.74***	-2.62	-3.95**	-4.26**	-0.27	-1.97**
Lagged variables below	(0.65)	(-0.11)	(-2.91)	(-1.36)	(-2.46)	(-2.36)	(-0.18)	(-2.15)
$r_{t-1}^{10y,US}$	-7.91***	-1.74**	-2.19***	-1.59	-5.41***	-1.92*	-3.93***	-2.18***
	(-3.97)	(-2.08)	(-2.82)	(-1.22)	(-3.40)	(-1.72)	(-4.66)	(-5.39)
$r_{t-1}^{10y,Foreign}$	6.29***	1.75**	1.53**	2.06	3.54***	2.10**	3.26***	1.81***
	(3.94)	(2.38)	(2.40)	(1.57)	(3.10)	(2.50)	(4.24)	(5.13)
π_{t-1}^{US}	3.43*	3.63*	2.50	0.62	3.74	2.54	3.36**	4.13***
	(1.90)	(1.75)	(1.49)	(0.28)	(1.42)	(1.44)	(2.12)	(3.44)
$\pi_{t-1}^{Foreign}$	-3.35**	-7.17***	-4.24**	-1.57	-4.73*	-5.54**	-6.36***	-5.81***
	(-2.08)	(-3.24)	(-2.07)	(-0.68)	(-1.84)	(-2.63)	(-3.45)	(-4.37)
$Risk_{t-1}$	-0.02***	-0.00*	-0.00	-0.01*	-0.01**	-0.00	-0.01***	-0.00***
	(-3.46)	(-1.70)	(-0.56)	(-1.94)	(-2.51)	(-0.07)	(-3.30)	(-2.70)
η_{t-1}	1.34	1.22	-4.84**	-3.03*	-1.78	-3.23*	0.17	-0.42
	(0.68)	(0.63)	(-2.45)	(-1.77)	(-1.05)	(-1.84)	(0.25)	(-1.03)
N	99	99	98	99	99	99	99	692
F	98	62	42	14	36	33	31	43
Adjusted R^2	0.72	0.62	0.54	0.45	0.56	0.56	0.61	0.55 (within)
Shapley (Traditional)	55%	59%	74%	60%	50%	64%	69%	57%

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Q1 1999 to Q2 2023. Standard errors are Newey West standard errors with 4 lags. For the panel regressions, standard errors are Driscoll Kraay 1998 standard errors with 4 lags. Shapley reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables (Δr_t, Δr_t^*, π_t,π_t^*, q_{t-1}, $\frac{TB}{GDP_t}$ and its lags).

Table A9: Panel regression with foreign currency as the base currency

LHS: Δs_t	Panel for AUD	Panel for CAD	Panel for EUR	Panel for GBP	Panel for NZD	Panel for NOK	Panel for SEK	Panel for USD
$\Delta r_t^{10y,Home}$	-6.12***	-6.03***	-4.65***	-4.48***	-4.73***	-5.88***	-4.31***	-4.03***
	(-8.80)	(-7.63)	(-6.31)	(-4.08)	(-6.40)	(-9.58)	(-5.69)	(-4.86)
$\Delta r_{i,t}^{10y,Foreign}$	5.87***	4.69***	5.59***	4.07***	5.80***	5.16***	4.83***	4.88***
	(7.46)	(6.25)	(8.12)	(5.20)	(8.40)	(9.17)	(7.49)	(5.74)
π_t^{Home}	-5.88***	-5.68***	-4.07***	-4.05***	-4.51***	-5.81***	-4.27***	-2.77***
	(-8.58)	(-7.22)	(-5.30)	(-3.79)	(-6.03)	(-7.74)	(-5.73)	(-3.36)
$\pi_{i,t}^{Foreign}$	5.80***	4.38***	5.33***	3.95***	5.82***	4.65***	4.50***	4.45***
	(6.93)	(5.95)	(7.33)	(5.19)	(8.21)	(7.56)	(7.16)	(5.13)
$q_{i,t-1}$	-0.02***	-0.02***	-0.02***	-0.01**	-0.02***	-0.01***	-0.02***	-0.02***
	(-3.20)	(-3.10)	(-3.17)	(-2.12)	(-4.02)	(-2.92)	(-3.72)	(-2.64)
$\Delta\eta_{i,t}$	-1.17**	-1.46***	-1.19**	-1.02*	-1.88***	-1.05***	-1.09*	-1.86***
Lagged variables below	(-2.42)	(-2.90)	(-2.04)	(-1.96)	(-3.34)	(-2.66)	(-1.93)	(-2.85)
$r_{t-1}^{10y,Home}$	-0.33***	-0.26*	-0.25*	-0.31**	-0.18*	-0.34**	-0.38***	-0.49***
	(-2.90)	(-1.82)	(-1.90)	(-2.10)	(-1.67)	(-2.35)	(-3.49)	(-3.04)
$r_{i,t-1}^{10y,Foreign}$	0.43***	0.23*	0.30**	0.28*	0.34***	0.24*	0.32***	0.45***
	(3.83)	(1.78)	(2.13)	(1.93)	(3.03)	(1.65)	(3.01)	(3.31)
π_{t-1}^{Home}	5.63***	5.27***	3.82***	3.79***	4.38***	5.62***	3.92***	2.01**
	(7.89)	(6.40)	(4.96)	(3.46)	(5.56)	(7.34)	(5.08)	(2.42)
$\pi_{i,t-1}^{Foreign}$	-5.45***	-4.12***	-5.13***	-3.72***	-5.55***	-4.44***	-4.17***	-3.97***
	(-6.23)	(-5.26)	(-7.17)	(-4.74)	(-7.55)	(-7.13)	(-6.25)	(-4.41)
$\eta_{i,t-1}$	-0.22	-0.23	-0.35	-0.03	-0.07	-0.09	-0.34**	-0.60***
	(-1.19)	(-0.85)	(-1.37)	(-0.12)	(-0.34)	(-0.47)	(-2.23)	(-3.33)
N	2087	2083	2081	2083	2087	2087	2071	2223
F	12.93	11.92	11.51	4.65	11.83	18.01	12.56	7.68
Adjusted R^2	0.16	0.14	0.14	0.09	0.14	0.15	0.11	0.18
Shapley (Traditional)	96%	93%	94%	91%	78%	94%	88%	87%
Shapley (Financial)	4%	7%	6%	9%	22%	6%	12%	13%

Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023. In each panel regression, the column currency is treated as the home currency and U.S dollar is excluded as a foreign currency in this analysis. Standard errors are Driscoll Kraay 1998 standard errors with 12 lags. Shapley reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables (Δr_t, Δr_t^*, π_t,π_t^*, q_{t-1}, and its lags).

Figure A1 cumulative fitted exchange rate value generated, quarterly frequency model in Table 2

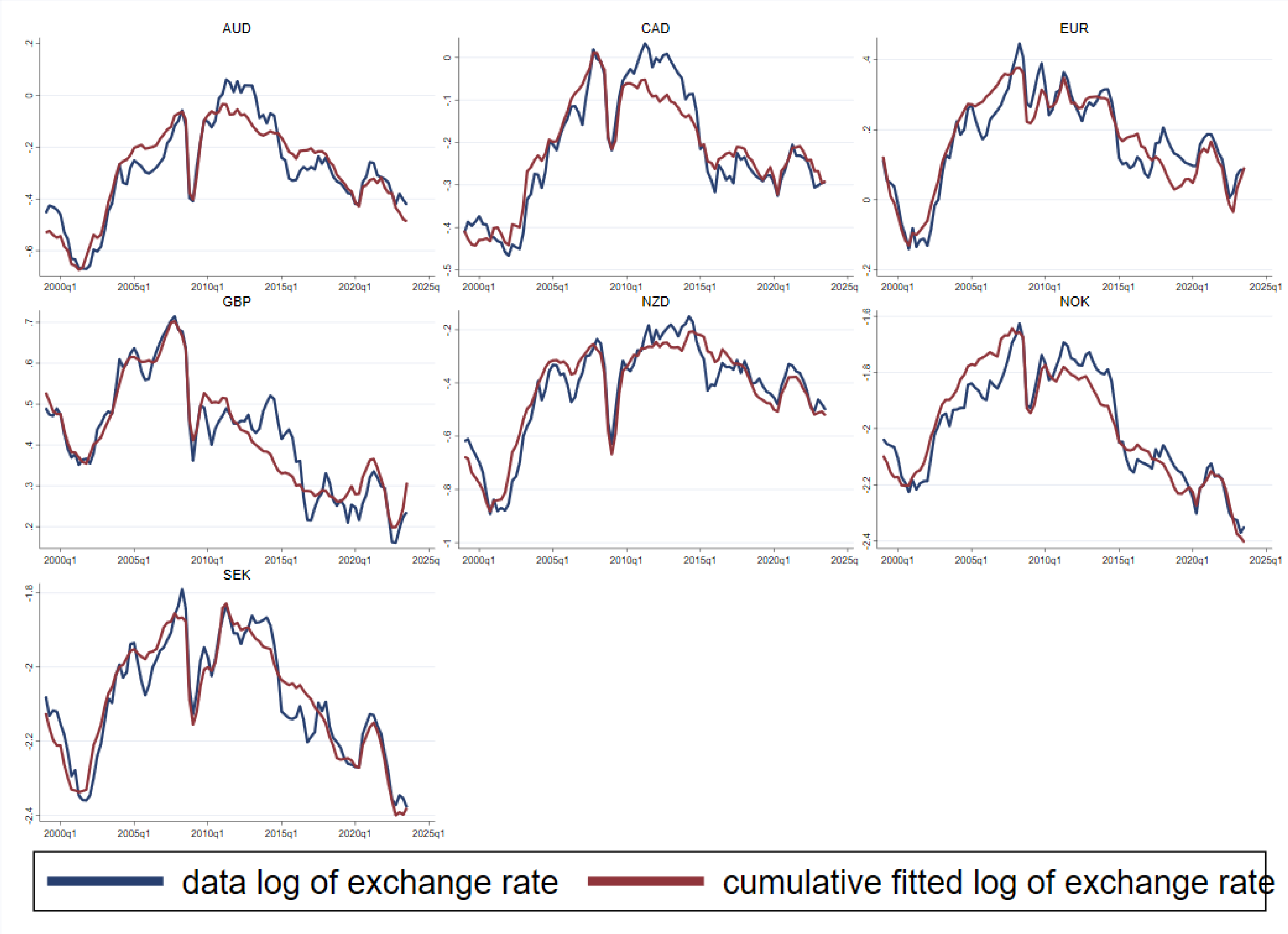


Figure A2 cumulative fitted exchange rate value generated, monthly frequency model in Table 2

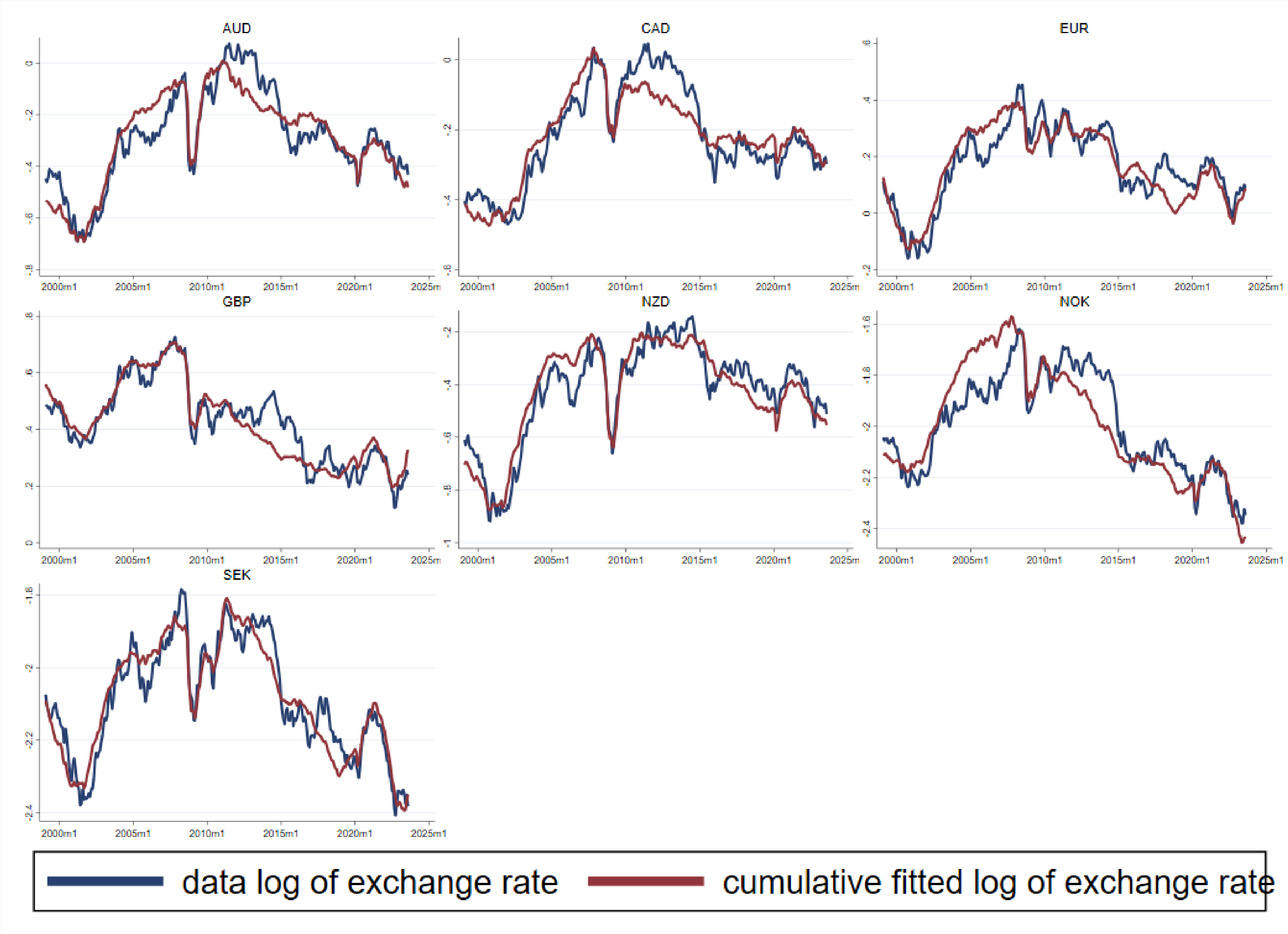
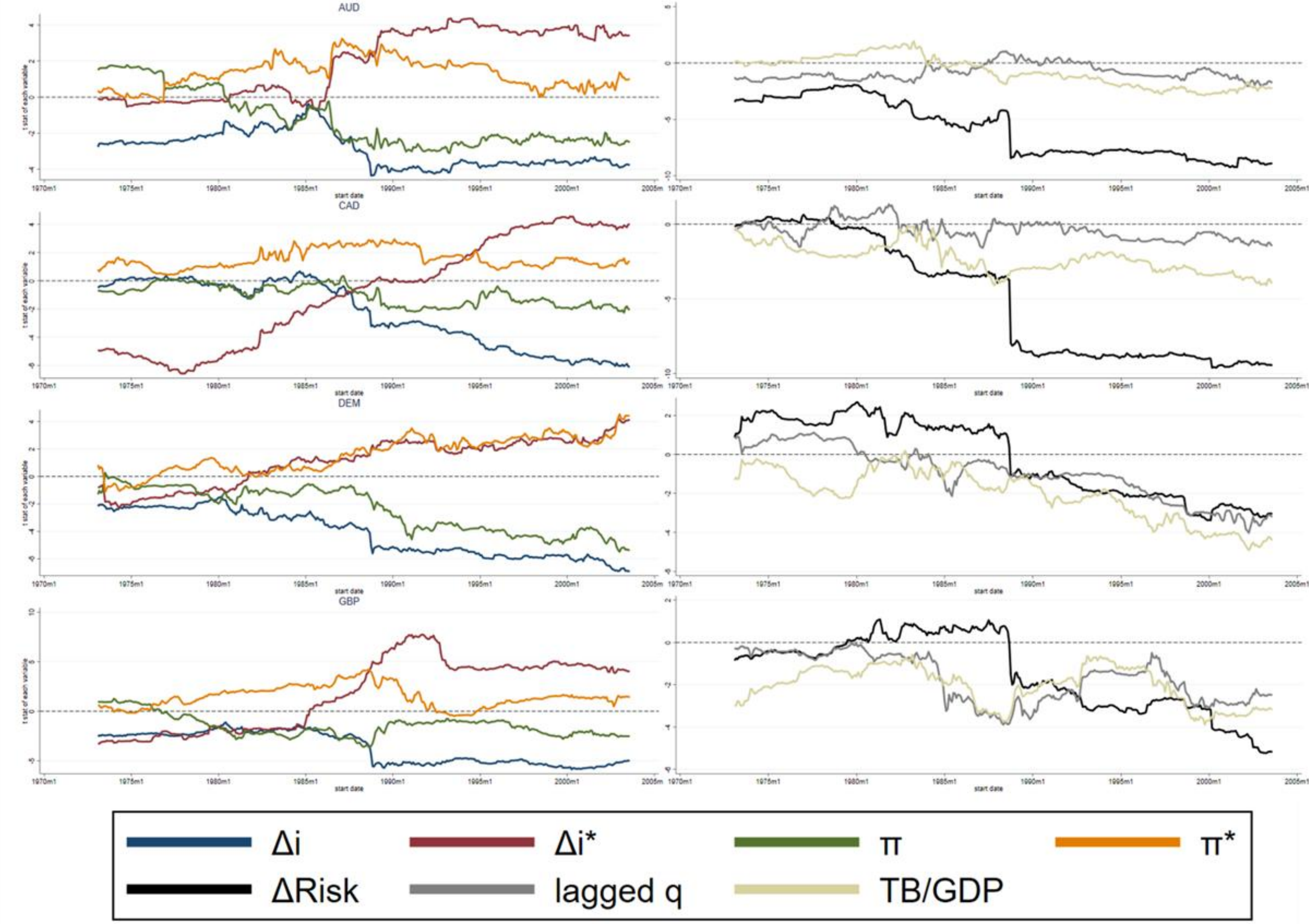
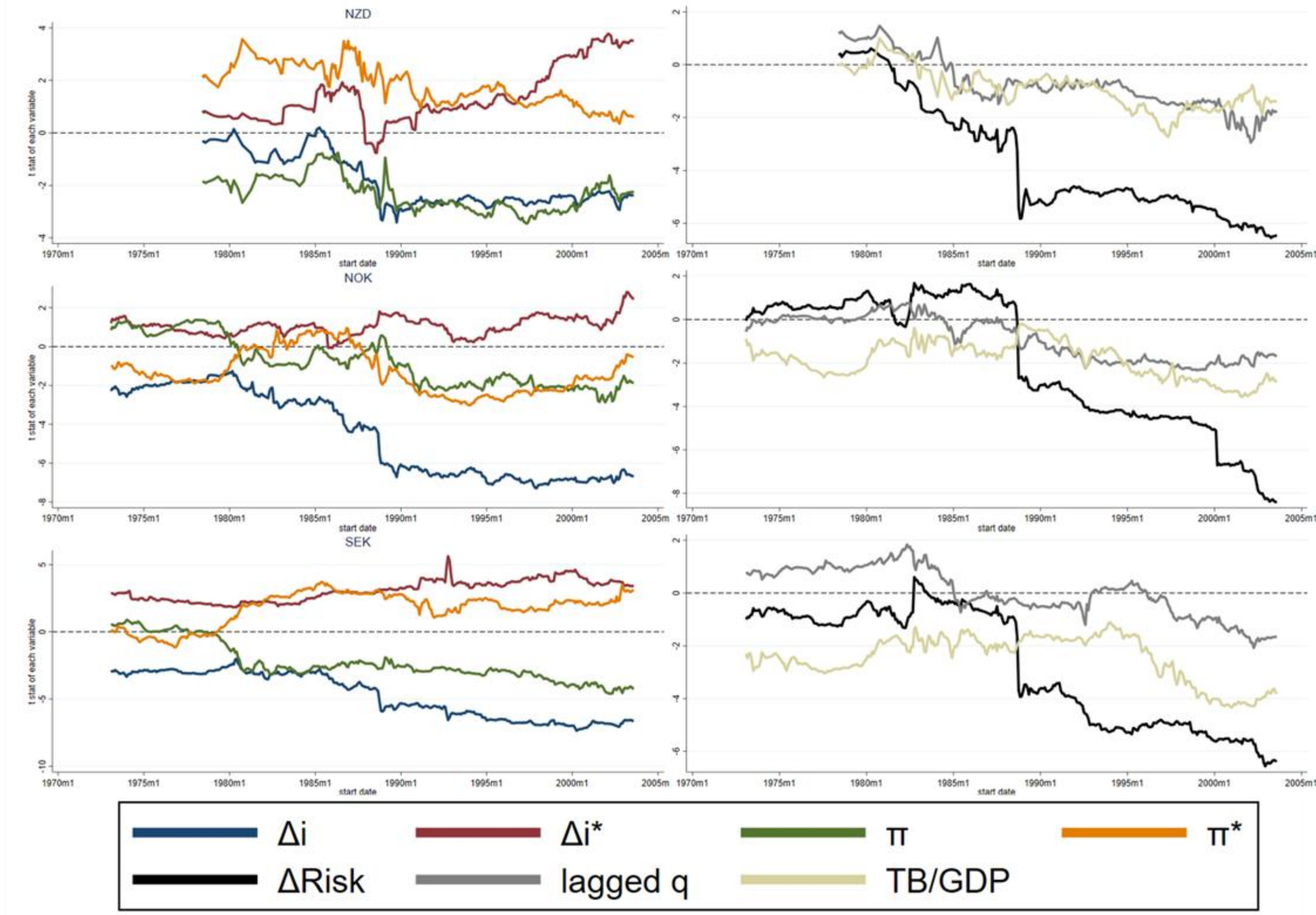


Figure A3: t statistics of 20-year rolling window regressions of equation (1)



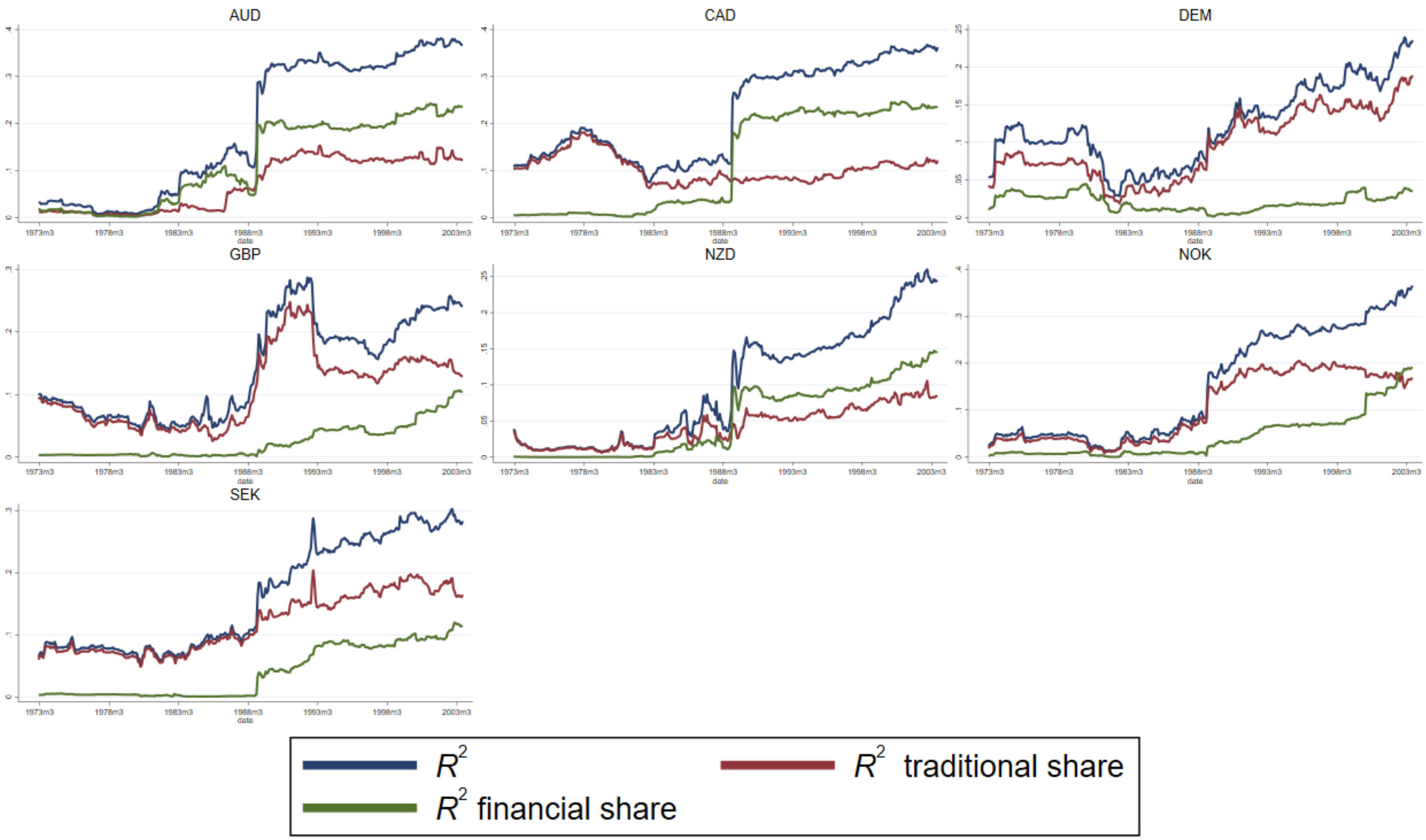
Note: The figure reports the t-statistics of each of the variables in equation (1) with a 20-year rolling window regression. X-axis correspond to the start date of the rolling window regression . Δr_t and Δr_t are the change of home and foreign real interest rate, π_t and π_t^* are the home and foreign 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 A3 continued: *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.

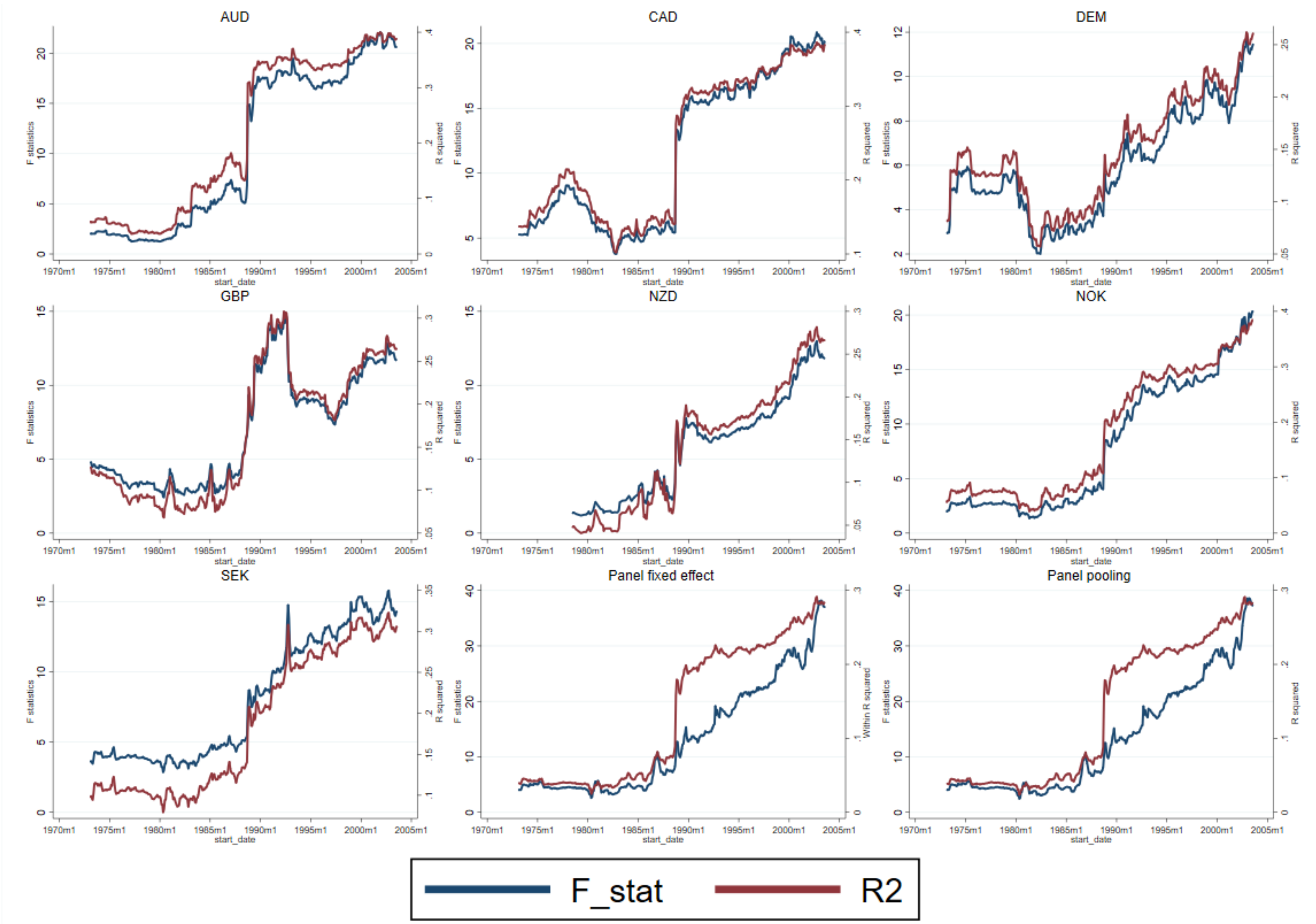
Figure A4 20-year rolling window R^2 by country and the Shapley contribution



Note: The figure reports the R^2 in equation (1) with a 20-year rolling window regression. Shapley (R^2 traditional) reports the Shapley value regression decomposition that computes the average marginal contribution of the traditional variables ($\Delta r_t, \Delta r_t^*, \pi_t, \pi_t^*, \frac{TB}{GDP_t}$). Shapley ($R^2$ financial) reports the same Shapley value decomposition for $\Delta RISK_t$. 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 A5: F-statistic and R2 of 20-year rolling window regressions of

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



Note: X-axis corresponds to the start date of the rolling window. The first regression is Mar 1973-Feb 1993. The last regression is Sep 2003-Aug 2023.

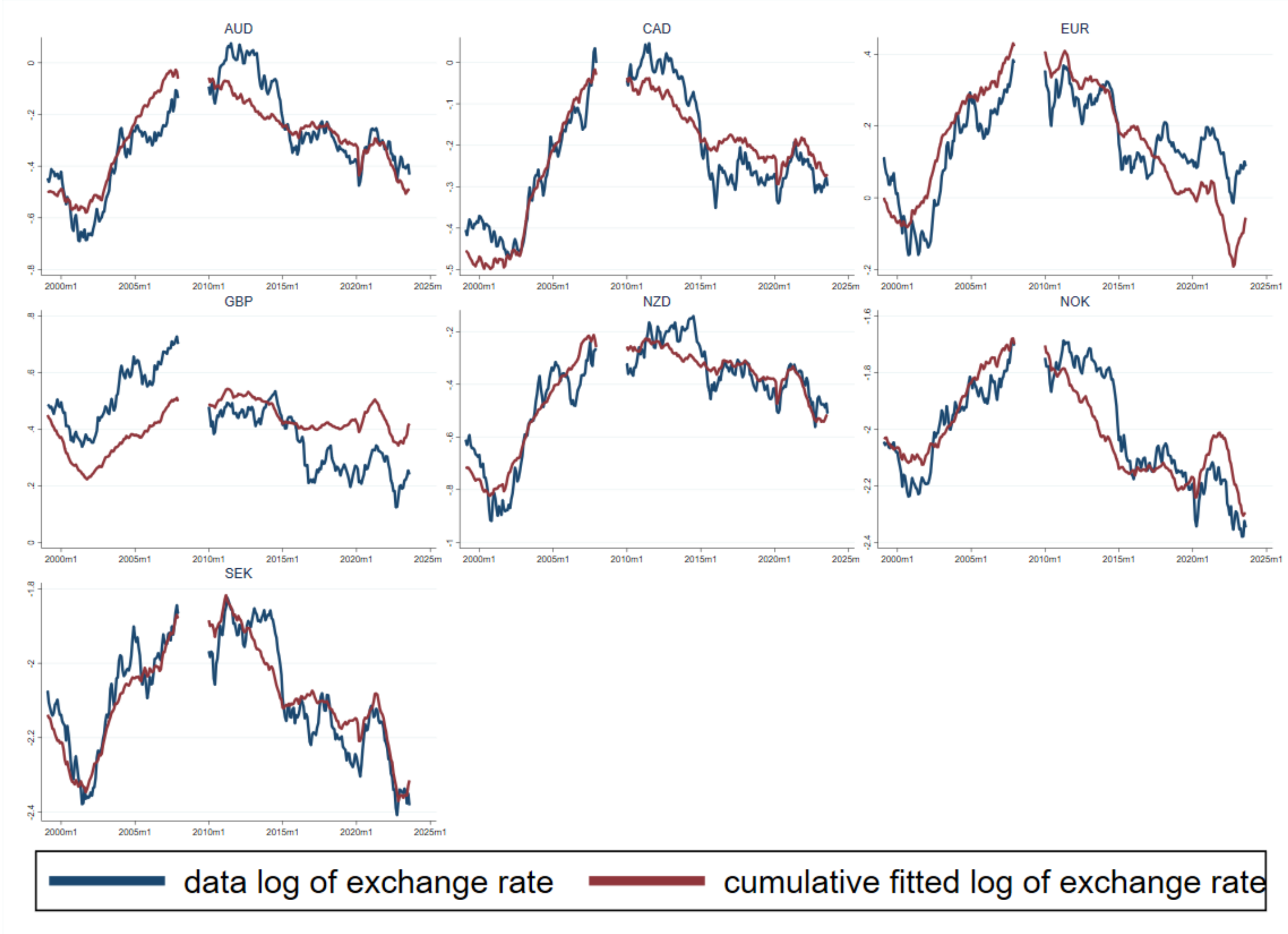
Appendix C: In sample regression analysis excluding 2008-2009

Table A10: Table 1 Baseline regression, excluding 2008-2009

	AUD	CAD	EUR	GBP	NZD	NOK	SEK	Panel fixed effect
Δr_t	-1.04***	-1.76***	-2.39***	-1.16***	-0.79**	-1.83***	-1.86***	-1.37***
	(-3.04)	(-6.29)	(-6.96)	(-4.63)	(-2.33)	(-4.75)	(-5.14)	(-5.22)
Δr_t^*	0.79***	1.17***	2.12***	1.11***	0.60	0.09	0.48	0.68***
	(3.24)	(4.48)	(5.84)	(3.30)	(1.58)	(0.44)	(1.11)	(3.60)
π_t	-0.21*	-0.16	-0.57***	-0.50***	-0.49***	-0.14	-0.61***	-0.33***
	(-2.11)	(-1.04)	(-4.39)	(-4.51)	(-3.49)	(-1.19)	(-6.23)	(-4.42)
π_t^*	0.03	0.12	0.47***	0.28**	0.29*	-0.27*	0.27***	0.16**
	(0.20)	(0.66)	(3.51)	(3.53)	(2.11)	(-1.77)	(3.80)	(2.89)
$\frac{TB}{GDP_t}$	-0.46***	-0.46***	-0.59***	-0.83***	-0.38*	-0.58***	-0.75***	-0.54***
	(-2.81)	(-4.73)	(-4.00)	(-3.91)	(-1.71)	(-3.16)	(-4.12)	(-4.34)
$\Delta RISK_t$	-0.03***	-0.02***	-0.01*	-0.01***	-0.03***	-0.03***	-0.02***	-0.02***
	(-6.65)	(-8.21)	(-1.66)	(-2.14)	(-5.49)	(-4.12)	(-3.60)	(-6.28)
$\Delta\eta_t$	-1.49	-2.51***	-1.95*	-1.77*	-1.59**	-1.32	-2.02**	-1.58**
	(-1.68)	(-3.18)	(-1.91)	(-2.29)	(-1.94)	(-1.54)	(-2.13)	(-2.88)
q_{t-1}	-0.01	-0.01	-0.02*	-0.03**	-0.01	-0.03***	-0.00	-0.01*
	(-1.08)	(-1.27)	(-1.99)	(-2.99)	(-1.46)	(-2.63)	(-0.60)	(-1.76)
N	272	272	271	272	272	272	272	1903
F	11.96	13.09	10.34	5.89	6.88	10.68	8.97	15.45
Adj. R^2	0.24	0.26	0.22	0.13	0.15	0.22	0.19	0.17 (within)
Shapley (Trad)	28%	43%	88%	68%	30%	55%	71%	51%
Shapley (Fina.)	70%	55%	8%	24%	67%	40%	29%	47%

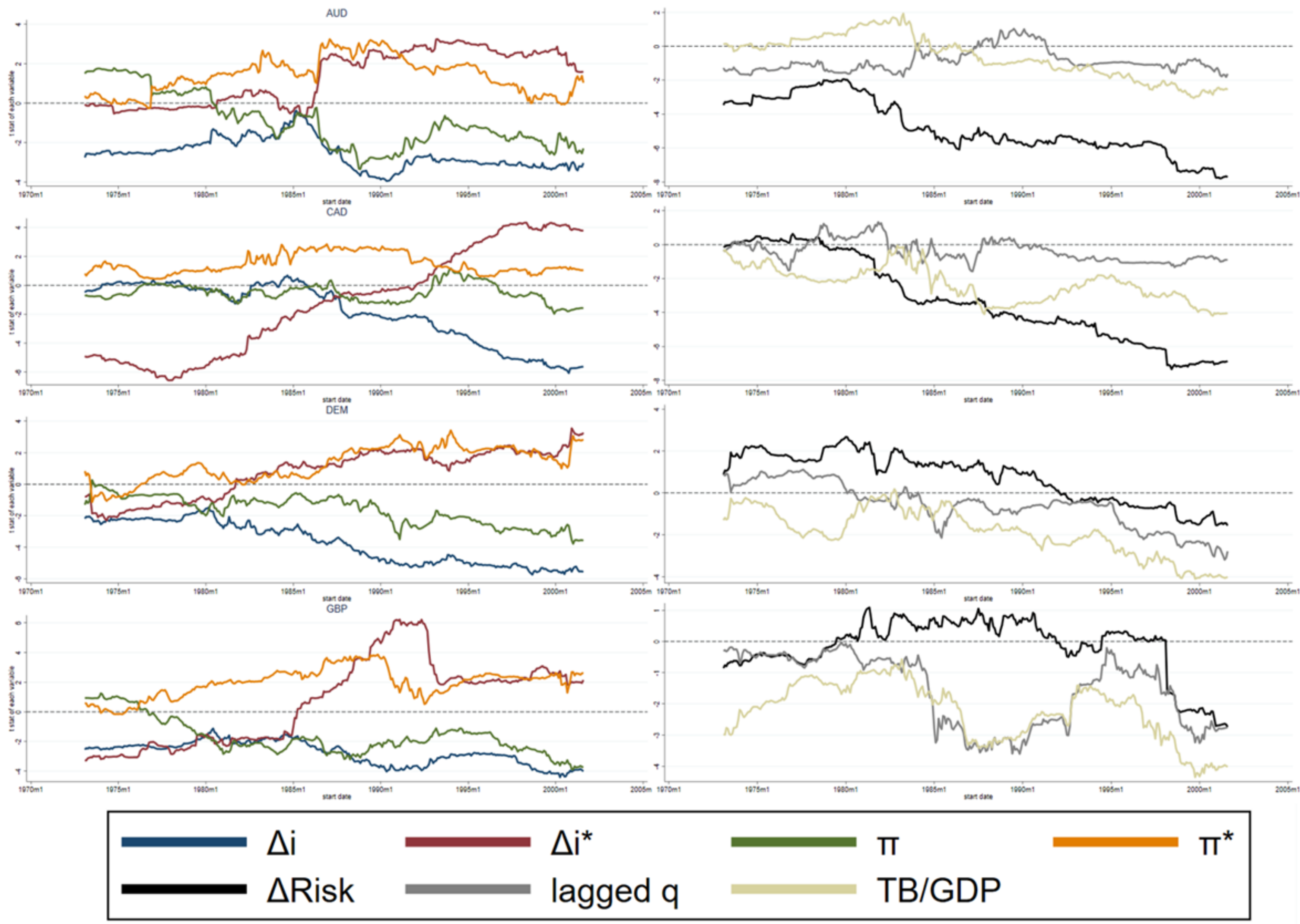
Note: t-statistics in parentheses. * p<0.1, ** p<0.05, *** p<0.01. Sample period is from Jan 1999 to Aug 2023 but excludes 2008-2009. The explanatory variable in all regression is the change of U.S. exchange rate with the currency in the column head. Standard errors are Newey West standard errors with 12 lags. 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).

Figure A6: Data and model implied exchange rates (excluding 2008-2009)



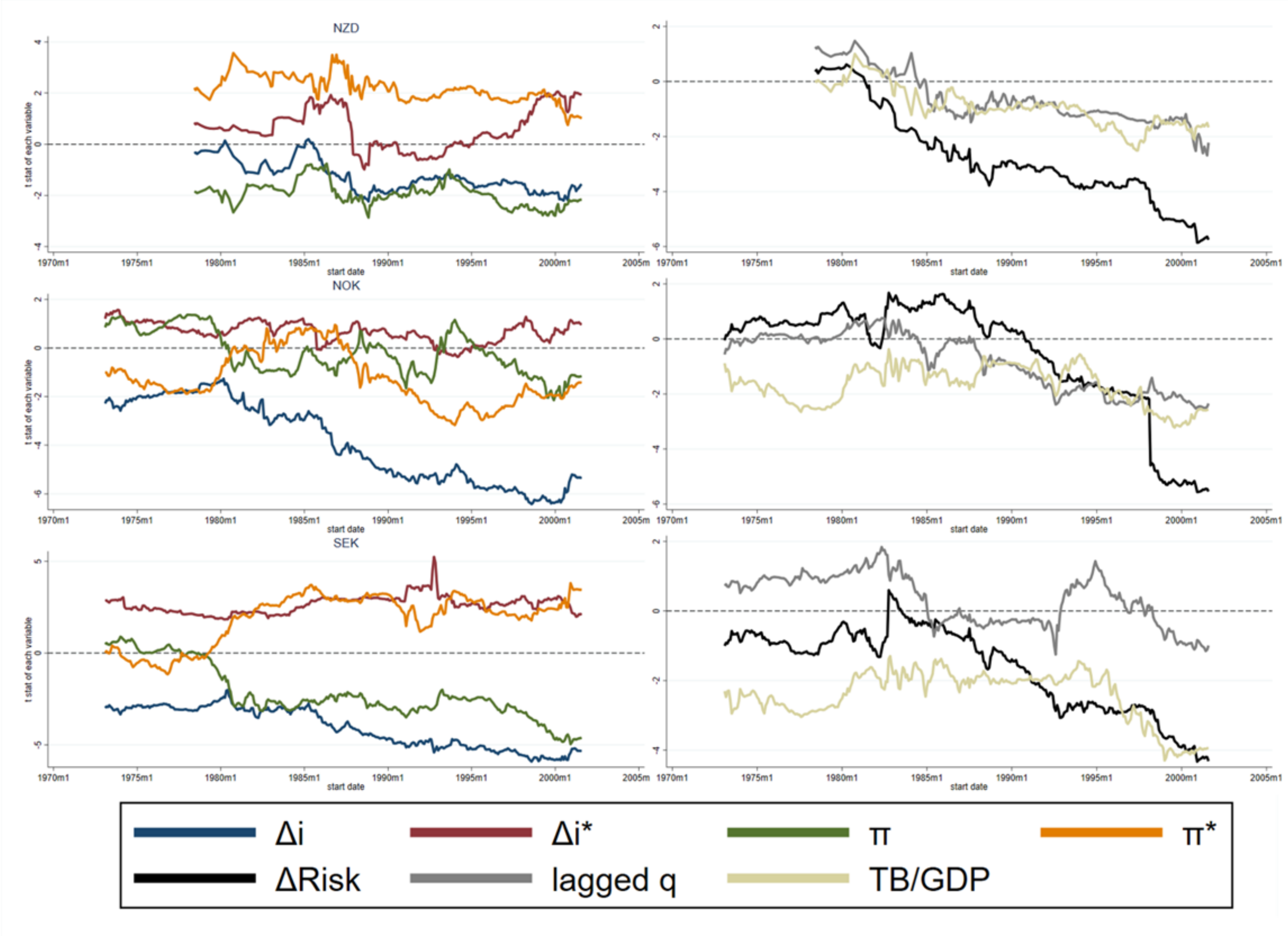
Note: The figure reports the log of exchange rate of each currency and the cumulative sum of model implied exchange rate change. The sample period is from Jan 1999 to Aug 2023 and excludes 2008-09. The mean value of the model implied log level of exchange rate is adjusted to have the same mean as the 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 A7: *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 A8: *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 A9: R^2 and Shapley contribution of 20-year rolling window regressions of equation (1), excluding 2008-2009

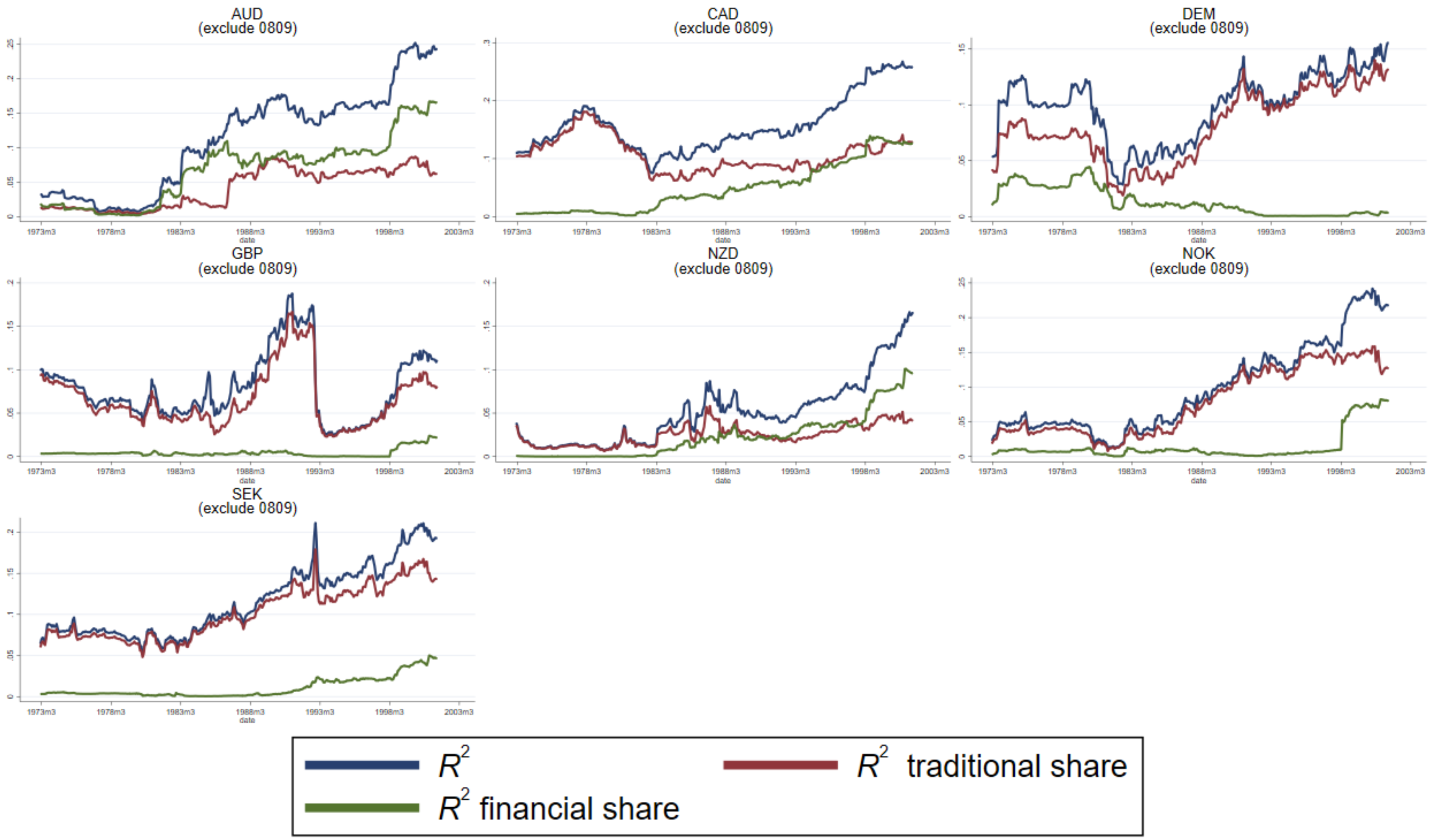
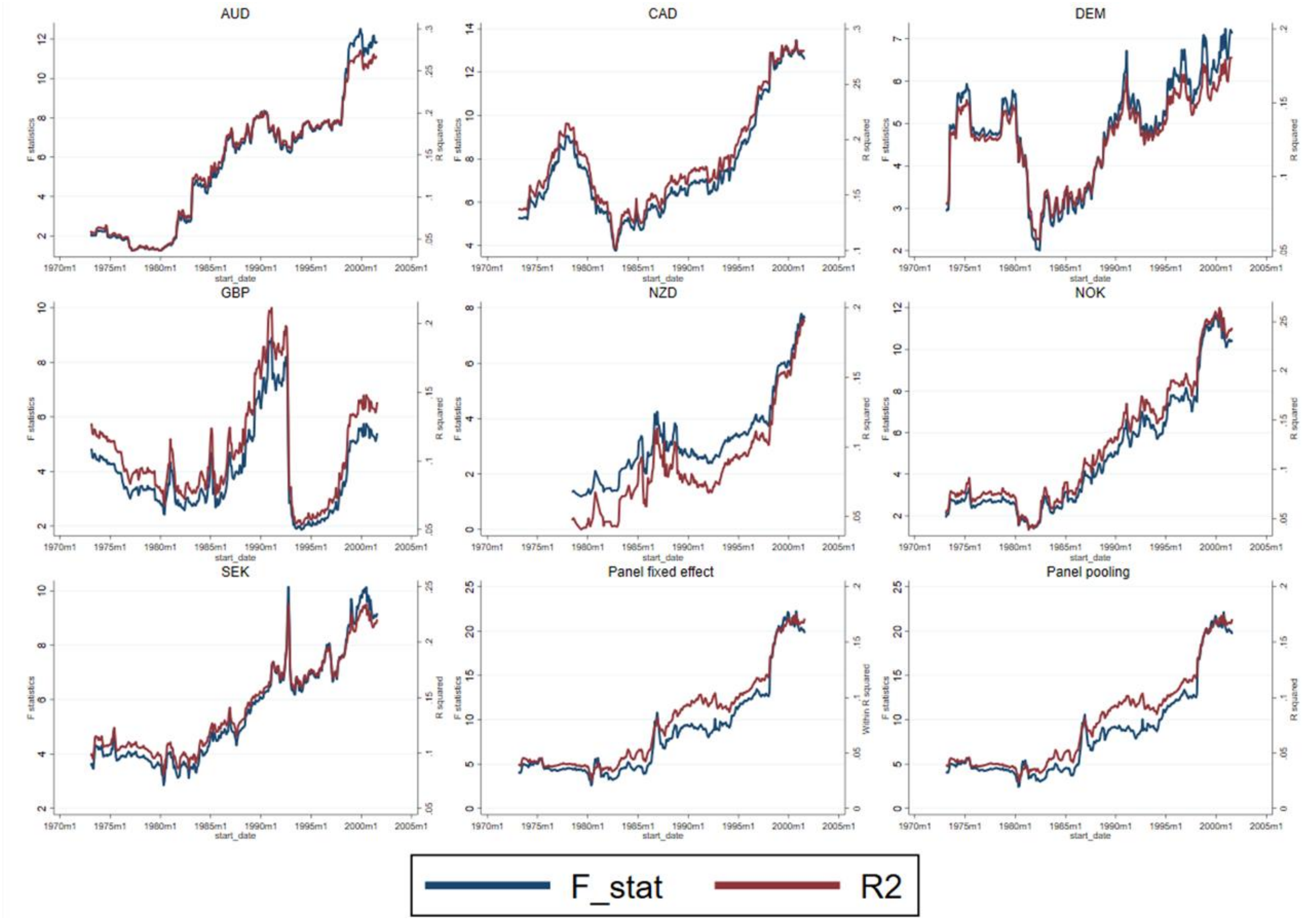


Figure A10: F-statistic and R^2 of 20-year rolling window regressions of equation (1), excluding 2008-2009



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.

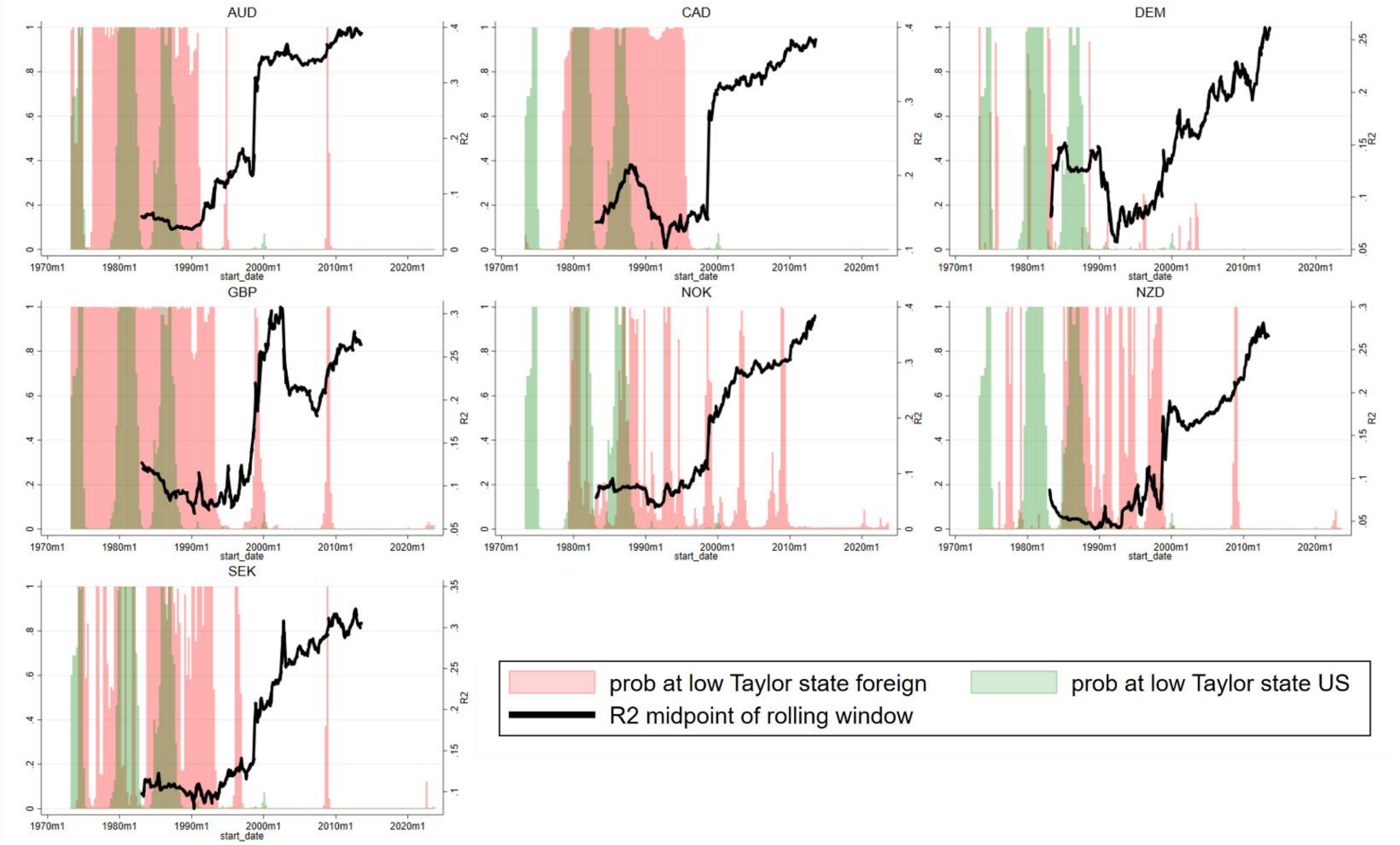
Appendix D: Robustness and additional evidence for section 4 on inflation credibility channel

Robustness and additional tables of Exercise 1: Difference in difference and Taylor rule estimation

Alternative Markov switching Taylor rule estimation

In the main text, the Markov switching model estimated impose a probability constraint such that the high Taylor coefficient state is an absorbing state. This corresponds to the case once a country gain inflation credibility, she always maintains the status. In the figure below, we report the same exercise but relax the probability constraint. The main message remains intact.

Figure A11 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.

Table A11: Two-year regression of exchange rate and interest rate before and after inflation targeting.

Regression: $\Delta s_{i,t} = \alpha + \beta_1 \Delta r_t + \beta_2 \Delta r_{i,t}^* + \beta_3 \pi_t + \beta_4 \pi_{i,t}^* + \beta_5 q_{i,t-1} + u_{i,t}$

Base currency	Panel for AUD	Panel for CAD	Panel for DEM	Panel for GBP	Panel for NZD	Panel for NOK	Panel for SEK
β_1, six to four years prior to the cutoff	-0.06	-0.02	0.44	2.01***	-0.28	0.45*	-0.16
	(0.19)	(0.39)	(0.66)	(0.53)	(0.20)	(0.24)	(0.10)
β_1, four to two years prior to the cutoff	0.93**	-0.48	0.30	0.86	-0.37	-0.14	-0.19
	(0.40)	(1.13)	(1.22)	(0.68)	(0.36)	(0.46)	(0.15)
β_1, two-year sample before the cutoff	-1.04	-0.62	0.08	-0.49	0.30	-0.30	-0.36
	(0.79)	(0.86)	(0.53)	(0.71)	(0.72)	(0.25)	(0.22)
β_1, two-year sample after the cutoff	-1.35***	-0.61**	-1.47**	-2.98***	-0.71**	-0.59***	-0.46**
	(0.33)	(0.24)	(0.56)	(1.03)	(0.26)	(0.21)	(0.22)
Cutoff date	1989 Dec	1991 Jan	1990 Jan	1992 Oct	1989 Dec	1992 Mar	1993 Jan

Note: The table reports the coefficients on home interest rate for each column head currency as the home currency. Driscoll Kraay (1998) standard errors with 4 lags in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Observation for each panel regression is 168.

Table A12: Regressions with interaction with inflation targeting treatment dummy variables.

Panel fixed effect regression			
Δr_t	-0.707***	$IT_{i,t} \times \Delta r_t$	-1.474***
	(-4.796)		(-7.642)
Δr_t^*	0.026	$IT_{i,t} \times \Delta r_t^*$	0.805***
	(0.345)		(6.769)
π_t	-0.037	$IT_{i,t} \times \pi_t$	-0.271***
	(-1.096)		(-4.041)
π_t^*	0.008	$IT_{i,t} \times \pi_t^*$	0.191***
	(0.379)		(3.733)
$\Delta RISK_t$	-0.001	$IT_{i,t} \times \Delta RISK_t$	-0.020***
	(-0.698)		(-6.851)
q_{t-1}	-0.005	$IT_{i,t} \times q_{t-1}$	-0.005
	(-1.068)		(-1.200)
TB/GDP_t	-0.094	$IT_{i,t} \times TB/GDP_t$	-0.168**
	(-1.064)		(-2.386)
		$IT_{i,t}$	-0.002
			(-0.906)
N	5337	Joint F test statistics of all the $IT_{i,t}$ interacted terms are zero (Chow test statistics)	12.46***

Note: Driscoll Kraay (1998) standard errors with 12 lags in parentheses: * p<0.1, ** p<0.05, *** p<0.01.

Table A13: Structural break selection: Inflation policy dates vs 2008

	Inflation policy cutoff dates	2008 Jan
Joint F test statistics of all the $IT_{i,t}$ interacted terms are zero (Chow test statistics)	12.46***	8.66***
Log-likelihood	12570	12547
Akaike Information Criterion (AIC)	-25092	-25046
Bayesian Information Criterion (BIC)	-24927	-24881

Robustness and additional tables of Exercise 2: Instrumental variable regressions

Table A14: Instrumental variable regressions with different standard errors and joint instrument

	Robustness: Table 9 (standard errors adjusted for 3 lags)	**Alternative Instrument: lagged change of each other country's CBI (standard errors adjusted for 3 lags)**
$\Delta\widehat{CBI}_{i,t}$	2.03***	2.18***
	(2.67)	(6.26)
First stage F statistic	16	37
Hansen J statistic		4.71 (*p* value 0.45)
N	210	210

Note: *t*-statistics in parentheses: * p<0.1, ** p<0.05, *** p<0.01. Sample period is from 1973 to 2023. The left-hand side variable in the regression is the annual change of R-squared from the exchange rate regression. The right-hand side variable is the change of central bank index from Romelli (2025). Driscoll Kraay (1998) standard error with 3 lags is reported.

Robustness and additional tables of Exercise 3: High frequency regressions

Table A15: First stage regressions of the instrumental variable regression of Table 10

LHS variable:	$\Delta OIS_{US,t}$	$\Delta OIS_{i,t}$
	(1)	(2)
U.S. inflation surprise	5.72***	1.62**
	(2.60)	(2.24)
Foreign inflation surprise	0.29	3.05***
	(1.08)	(6.76)
$\Delta risk_t$	-0.13	-0.12
	(-0.49)	(-1.43)
$\Delta \eta_{i,t}$	0.07	0.04*
	(1.2)	(1.87)
N	2354	2354

Appendix E: Model simulation

We simulate a standard 2-country New Keynesian model. First, we show that when the Taylor condition is satisfied, the artificial data from the model will reproduce regression results very similar to those we report in the body of the paper using actual data. Then we show that when the Taylor condition is not satisfied, the regressions fit poorly, with low R^2 values, low t-statistics, and "incorrect" signs for some of the parameters.

The model consists of a household sector, a firm sector that produces output using labor, a monetary policy rule, and a global financial intermediary that intermediates all global capital flows. The model is very similar to that of Itskhoki and Mukhin (2021) and Cormun and De Leo (2026), except that there are no noise traders. The model is also very similar to that of Kekre and Lenel (2024b), with the embellishment of slow price adjustment by Calvo pricing, as in New Keynesian models.

Because the equations of the model are so familiar, we mostly present first-order conditions for households and firms, and equilibrium conditions, all log-linearized.

Households in each country consume home and foreign aggregates. The two aggregates are imperfect substitutes, and we assume home bias in preferences. They get disutility from work, which they supply to firms in their own country.

Households can only trade bonds, and only with the global financial intermediary. Hence, home households can borrow or lend with home-currency denominated nominal bonds, and foreign households borrow or lend with foreign-currency denominated bonds. The financial intermediary takes the other side of transactions with households.

Firms in each country produce output of their good using only labor. The underlying assumption is standard in New Keynesian models – each firm is a monopolist in the production of a unique good, and hence has pricing power, and all goods produced within a country are imperfect substitutes for each other. We assume producer-currency pricing (PCP), as in the papers cited above. This means firms in each country set only one price, in their own currency, and the law of one price holds for internationally traded goods. But real exchange rates are not constant because consumers have home bias in consumption, so consume different baskets at home and abroad.

Financial intermediaries trade bonds with households in each country. They have a one-period horizon, exactly as in Itskhoki and Mukhin (2021) and Kekre and Lenel (2024b). Their capacity for intermediation is not unlimited because they are risk averse. As in these two papers, when a country borrows more, the financial intermediary must lend to that country and borrow from the other. This leads the risk-averse intermediary to require a risk premium on the borrowing country's debt. (We assume that the risk premium is related to the level of debt, but is imperfectly measured, so there is a component that is unobservable to the econometrician.) In addition, the intermediary has a convenience shock as in Jiang et al. (2024) and Kekre and Lenel (2024a). An increase in the "convenience" of the home bond means the intermediary is willing to hold more of it for a given rate of expected return.

Monetary policy in each country sets a Taylor rule that targets its own country's CPI inflation rate, with nominal interest-rate smoothing.

There are eight exogenous shocks in total: disturbances to the home and foreign monetary policy rules; shocks to home and foreign aggregate demand; shocks to home and foreign markups in the price-setting rules; shocks to relative convenience; and shocks to the unobserved measurement error of the bond risk premium. We suppress shocks to TFP, since in our simulations, reasonably parameterized processes seem to have little effect on the regression results.

Here we provide some intuition of the signs of the regression coefficients for real interest rates, expected inflation, the trade deficit and the convenience yield. Any regression coefficient is interpreted as the long-run relation between an independent variable and the dependent variable, holding the other independent variables constant. For example, a negative coefficient in the regression of the change in the exchange rate on the change in the home real interest rate means that on average, holding the other variables constant, when the home real interest rate change is positive, the change in the exchange rate is negative (home appreciation.) In other words, the regression coefficients are not impulse responses (if x goes up at time t what happens to y at time t?) because the exogenous variables are persistent and there is slow adjustment of the economic system.[23]

Demand shocks, monetary policy shocks and cost-push shocks work primarily through their effect on real interest rates and expected inflation, while unobserved risk premium shocks and convenience shocks work through the coefficients on debt accumulation and the measure of convenience yield. In fact, the accumulation of debt can be influenced by all the exogenous shocks, but as we note below is associated with a weakening of the currency.

Demand shocks affect exchange rates very much as in Kekre and Lenel (2024b), and indeed as they would in classical models. Relative to these studies, in the New Keynesian model, real interest rates are mediated through monetary policy. The path of real interest rates is determined by the monetary policy rule – how the current nominal interest rate reacts to inflation and how adjustment occurs over time.

Consider a growing home demand, first if prices adjusted flexibly. Increasing demand implies lower saving and more borrowing by the home country. This tends to raise the real interest rate of the home country. The demand disturbances are temporary (but persistent), so the real exchange rate is expected to converge to its unconditional mean. When real interest rates in the home country are increasing, there must be an appreciation in the current real exchange rate to maintain uncovered interest parity (holding risk premia constant, as controlled in the regression)

When prices are sticky and there is a monetary policy rule, the dynamics of adjustment are influenced by the monetary policy reaction. Higher demand for home goods is associated with

[23] When we say an increase in a variable is, for example, associated with an appreciation of the home currency, we mean the change in the exchange rate between t-1 and t is associated with the level or change of the time t variable. In other words, we are not making forecasts of the exchange rate between t and t+1. "Disconnect" is concerned primarily with contemporaneous correlation of the exchange rate with macroeconomic variables. In fact, as Engel and West (2005) demonstrate, a wide class of models that imply contemporaneous correlation still imply little forecasting power.

higher home inflation. If the Taylor condition is satisfied, then real interest rates gradually increase in the home country, leading to a real appreciation as in the previous paragraph.

Monetary policy disturbances and cost-push shocks are features introduced in the New Keynesian framework, but their effect on the relationship between real interest rates and the exchange rate are similar. When monetary policy is tightening, real interest rates rise and the currency appreciates. Cost-push shocks lead to higher inflation, and under a stable Taylor rule that leads to higher real interest rates and an appreciating currency.

As noted in the main body of the paper, because of slow adjustment (arising from slow adjustment of prices in the Calvo pricing framework, interest rate smoothing in the Taylor rule, and gradual accumulation of assets), the current real interest rate does not completely reflect the stance of monetary policy. Holding the change in the current real interest rate constant, higher inflation is associated with future increases in real interest rates under a stable Taylor rule. In turn, the currency appreciates.

The convenience yield is an exogenous variable here, and an increase in the home convenience yield reduces the expected return required on the home bond relative to the foreign bond. Intuitively, this can be understood as an increase in a non-pecuniary yield to the home bond and is associated with an appreciation of the home currency.

Finally, various exogenous variables may work to increase or decrease the trade balance, or the rate of borrowing/lending of either economy. Because the financial intermediary is risk averse, as the home country borrows more (has a higher deficit), the intermediary's holding of home debt increases requiring a greater risk premium on that debt, which is associated with a depreciation of the home currency. The sign of the coefficient here may in general depend on the strength of the portfolio balance effect for the intermediary.

We calibrate the model using standard parameter values established in the literature at monthly frequency. Specifically, the discount factor (β) is set to 0.995. We adopt log utility for consumption ($\sigma = 1$), set the Frisch labor supply elasticity (φ) to 1, the trade elasticity (η) to 1, and the home bias parameter ($1 - \nu$) to 0.9. The financial intermediation friction (γ) is specified as 0.005. Finally, we set the price stickiness parameter (θ) to 0.99, the Taylor rule inflation response coefficient (ϕ_π) to 1.01 (ensuring determinacy), and the interest rate smoothing parameter (ρ_R) to 0.93. The standard deviation of (home and foreign) markup shocks and demand shocks are set at 0.002, 0.001. The standard deviation of (home and foreign) monetary shock is 0.001. The standard deviation of convenience yield/risk premium shock is 0.0005. We allow for some persistence for demand shock and risk premium shock, to be at 0.9. When we introduce a sunspot shock, its standard deviation is set the same as the monetary shocks, 0.001.

In Table A16, we report the exchange rate model regression using simulated data. The model is simulated for 280 periods (matching the monthly sample from 2000–2023) and run 1000 times. The regression regresses the monthly change in the nominal exchange rate on changes in the home real interest rate (nominal minus model-expected inflation), foreign real interest rate, model home and foreign expected inflation, changes in the convenience yield/risk premium, and the trade balance-to-output ratio.

In the first column, the coefficients are all statistically significant and closely resemble the results in Table 1 of the paper. The size of the unobserved UIP shock is set to achieve a realistic R^2 of about 40%. We keep the same size of the unobserved UIP shock when we change the monetary policy credibility.

Table A16: Regression with simulated data

	Taylor principle is satisfied	Taylor principle is not satisfied sunspot s.d. = 0	Taylor principle is not satisfied sunspot s.d. = 0.001
	(1)	(2)	(3)
$\Delta(r)_t$	-8.79***	0.59	0.23
$\Delta(r^*)_t$	8.84***	-0.57	-0.25
$(\pi)_t$	-2.31**	4.79***	4.40***
$(\pi^*)_t$	2.28**	-4.81***	-4.38***
$\Delta\omega_t$	-5.17***	-0.87	-0.80
$\Delta b_t/Y_t$	-2.76***	0.01	0.04
N	280 each, simulated 1000 times	280 each, simulated 1000 times	280 each, simulated 1000 times
Adjusted R^2	0.40	0.09	0.08

Note: t-statistics are reported. * p<0.1, ** p<0.05, *** p<0.01.

When the Taylor condition is not satisfied, we set ϕ_π to be 0.5. There are an infinite number of potential "stable" solutions to the system. One of the solutions implies that a complicated linear function of the variables in the system is completely predictable. We find other solutions using the sunspot structure of Bianchi and Nicolo (2021). We display regression results from artificial data generated from two different solutions to the model in columns (2) and (3) in Table A16. In the first, we take the case in which there is no sunspot (so that the linear combination of variables is predictable.) In the second case, we introduce a sunspot with a low standard deviation and set its standard deviation equal to the standard deviation of shock for the monetary policy shock (0.001). In both cases, we see that the R^2 dramatically dropped from 40% to 9%. We observed many incorrect signs among the estimated coefficients, such as a positive coefficient for home inflation. These findings are all consistent with the pattern observed in the "old days".

On the one hand, the solution with the sunspot no longer implies there is any linear combination of variables that is perfectly predictable. On the other hand, our point is to emphasize that it is not the noise from the sunspot that leads to the poor fit of the empirical model with the artificial data, because we see the fit is poor even in the case of no sunspot. It is the failure of the monetary policy makers to react sufficiently strongly to increases in inflation that lead to the poor fit of the model when the Taylor condition is not met.[24]

[24] In fact, if there were a sunspot shock that had a high variance, it would lead to a good fit of the exchange rate regression, because the dependent variables are endogenous and also driven by the sunspot – though many of the coefficient signs would be counterfactual. So, it is not noise from a sunspot *per se* that leads to the poor fit we see in the actual and simulated data.

The model equations are listed below:

Definition of Variables

c_t (c_t^*) – log of home (foreign) consumption

w_t (w_t^*) – log of home (foreign) nominal wage

$y_{h,t}, y_{f,t}$ - log of home demand for home (foreign) good

$y_{f,t}^*, y_{h,t}^*$- log of foreign demand for foreign (home) good

$p_{h,t}, p_{f,t}$ - log of home-currency price of home (foreign) good

$p_{f,t}^*, p_{h,t}^*$ - log of foreign-currency price of foreign (home) good

p_t (p_t^*) – log of home (foreign) consumer price level

$\pi_{hh,t}$ ($\pi_{ff,t}^*$) – home (foreign) inflation of home- (foreign-) produced goods

π_t (π_t^*) – home (foreign) consumer price inflation rate

n_t (n_t^*) – log of home (foreign) employment

y_t (y_t^*) – log of home (foreign) output

i_t (i_t^*) – home (foreign) nominal interest rate

s_t - log of nominal exchange rate (home currency price of foreign currency)

ω_t - deviation from uncovered interest parity

b_t - home net claims on foreigners (relative to GDP)

nx_t - home net exports (relative to GDP)

q_t - log of real exchange rate (foreign price level relative to home)

$sunspot_t$ ($sunspot_t^*$) – home (foreign) sunspot disturbance

ϵ_t^d (ϵ_t^{d*}) – home (foreign) demand disturbances

ϵ_t^{mk} (ϵ_t^{mk*}) – home (foreign) mark-up disturbances

ϵ_t^m (ϵ_t^{m*}) – home (foreign) disturbances to monetary policy rules

ρ_t - disturbance to convenience yield of home bonds

u_t - unobserved component of intermediary's risk premium on foreign bonds

Household block:

Euler equation

The households Euler equation are described as:

(10) $-\sigma c_t = i_t - \sigma E_t c_{t+1} - E_t \pi_{t+1} + \epsilon_t^d$

(11) $-\sigma c_t^* = i_t^* - \sigma E_t c_{t+1}^* - E_t \pi_{t+1}^* + \epsilon_t^{d^*}$

Wage equation

(12) $w_t - p_t = \varphi n_t + \sigma c_t$

(13) $w_t^* - p_t^* = \varphi n_t^* + \sigma c_t^*$

where η is the elasticity of labor.

Demand for goods

(14) $y_{h,t} = c_t - \eta(p_{h,t} - p_t)$

(15) $y_{f,t} = c_t - \eta\left(p_{f,t} - p_t\right)$

(16) $y_{h,t}^* = c_t^* - \eta(p_{h,t}^* - p_t^*)$

(17) $y_{f,t}^* = c_t^* - \eta(p_{f,t}^* - p_t^*)$

<u>**Firm block**</u>

Phillips Curve

(18) $\pi_{h,t} = \kappa_p\left(w_t - p_{h,t}\right) + \beta E_t \pi_{h,t+1} + \epsilon_t^{mk}$

(19) $\pi_{f,t}^* = \kappa_p\left(w_t^* - p_{f,t}^*\right) + \beta E_t \pi_{f,t+1}^* + \epsilon_t^{mk^*}$

Domestic production

(20) $a_t + n_t = y_t$

(21) $a_t^* + n_t^* = y_t^*$

LOOP

(22) $p_{h,t}^* = p_{h,t} - s_t$

(23) $p_{f,t} = p_{f,t}^* + s_t$

Definition of CPI

(24) $p_t = (1 - \nu)p_{h,t} + \nu p_{f,t}$

(25) $p_t^* = (1 - \nu)p_{f,t}^* + \nu p_{h,t}^*$

Definition of inflation

(26) $\pi_{h,t} = p_{h,t} - p_{h,t-1}$

(27) $\pi_t = p_t - p_{t-1}$

(28) $\pi_{f,t} = p_{f,t}^* - p_{f,t-1}^*$

(29) $\pi_t^* = p_t^* - p_{t-1}^*$

Goods market clearing

(30) $y_t = (1 - \nu)y_{h,t} + \nu(y_{h,t}^*)$

(31) $y_t^* = (1-\nu)y_{f,t} + \nu(y_{f,t}^*)$

Taylor rules

(32) $i_t = \rho_R i_{t-1} + (1-\rho_R)\phi_\pi \pi_t + \epsilon_t^m$

(33) $i_t^* = \rho_R i_{t-1}^* + (1-\rho_R)\phi_\pi \pi_t^* + \epsilon_t^{m*}$

Exchange rate determination (s is the nominal exchange rate, q is the real exchange rate)

(34) $s_{t+1} - s_t = i_t - i_t^* + \omega_t$

(35) $\omega_t = \gamma(b_t + u_t) + \rho_t$

ω_t is a UIP wedge, b_t is NFA position, u_t is unobserved noise premium, ρ_t is the convenience yield.

(36) $\beta b_t = nx_t + b_{t-1}$

(37) $nx_t = y_t - c_t + \nu(p_{h,t} - p_{f,t})$

(38) $q_t = s_t + p_t^* - p_t$

sunspot

(39) $\pi_t - E_{t-1}\pi_t = -sunspot_t + Gamma \times sunpot_{t-1} + sunspotshock_t$

(40) $\pi_t^* - E_{t-1}\pi_t^* = -sunspot_t^* + Gamma \times sunpot_{t-1}^* + sunspotshock_t^*$

This is the Bianchi Nicolo sunspot structure. Gamma = 2 when the model features sunspot

Exogenous processes:

Convenience yield/risk premium disturbances

(41) $\rho_t = 0.9\rho_{t-1} + u_t^{Risk}$

Demand shocks

(42) $\epsilon_t^d = 0.9\epsilon_{t-1}^d + u_t^d$

(43) $\epsilon_t^{d*} = 0.9\epsilon_{t-1}^{d*} + u_t^{d*}$

Mark up shocks

(44) $\epsilon_t^{mk} = 0.0\epsilon_{t-1}^{mk} + u_t^{mk}$

(45) $\epsilon_t^{mk*} = 0.0\epsilon_{t-1}^{mk*} + u_t^{mk*}$

Monetary shocks

(46) $\epsilon_t^m = 0.0\epsilon_{t-1}^m + u_t^m$

(47) $\epsilon_t^{m*} = 0.0\epsilon_{t-1}^{m*} + u_t^{m*}$

Unobserved UIP disturbances

(48) $u_t = 0.0u_{t-1} + u_t^{UIP}$

Appendix F: Macro disconnect in Lilley et al 2022

In this appendix, we re-examine the recent macro disconnect phenomenon (Lilley et al., 2022). Our findings suggest that jointly accounting for a broader set of fundamental drivers reveals a robust connection between exchange rates and macroeconomic variables.

Since Lilley et al. (LMNS) focus on linking financial risk variables to exchange rates post-2008, their main text omits macroeconomic variables. Four exhibits in their appendix directly discuss the relationship between exchange rates and macro fundamentals.

1) Table A.4 that involves in sample regressions.
2) Figure A.6 that involves in sample rolling window regressions.
3) Figure A.1 that involves scatter plot of exchange rate and macro variables and
4) Table A.5 that involves Meese Rogoff style out-of-sample fit regressions

We perform our baseline regression *directly using data from the replication file of LMNS* and show that a strong connection between exchange rate and macro variables can be observed.

1) Table A.4 of LMNS

Table A17 reproduces the in-sample regressions from LMNS. They conduct in-sample tests using several regressions motivated by the literature: an UIP regression, a Taylor-rule fundamental regression, a Backus–Smith consumption regression, and others. These are time-series regressions in which the exchange rate is constructed as the simple average of non-U.S. G10 exchange rates, and foreign variables are defined as simple averages across non-U.S. G10 countries.

Panel B of their table A.4 focuses on the post-2007 sample (column (1) to (5) of TableA17), which is the sample we find a stronger connection between macro fundamentals and exchange rates. However, all of their regressions exhibit low R^2 values. But at the very least, the quarter-over-quarter inflation variable is statistically significant with the expected sign.

We add a new Column (6) alongside this table, which presents our preferred regression specification. We exclude Japan and Switzerland when constructing the average exchange rate and foreign variables. We use year-over-year (YoY) inflation instead of quarter-over-quarter (QoQ) inflation, as we believe it better captures underlying inflation dynamics with less noise and is more closely the target of the Federal Reserve and other central banks. Our baseline specification includes the change in the U.S. real interest rate (defined as the nominal rate minus current YoY inflation), the change in the foreign real interest rate, U.S. YoY inflation, foreign YoY inflation, the change in the GZ spread as a proxy for risk premia, and the change in convenience yield.

Table A17. Reproduction of Lilley et al Table A.4 and our baseline regression

	Panel B: 2007Q1-2019Q2					
	(1) Flow	(2) UIP	(3) Backus Smith	(4) Monetary	(5) Taylor	(6)
US flows	0.85					
	(0.16)					
$i_{t-1}^{US} - \bar{\iota}_{t-1}$		1.3				
		(0.61)				
$\pi_t^{US} - \bar{\pi}_t$			2.5	2.8	2.9	
			(0.65)	(0.57)	(0.54)	
$\Delta c_t^{US} - \overline{\Delta c}_t$			-0.75			
			(1.28)			
$\Delta y_t^{US} - \overline{\Delta y}_t$				0.50		
				(1.17)		
$\tilde{y}_t^{US} - \bar{\tilde{y}}_t$					0.0063	
					(0.099)	
Δr_t^{US}						-2.91***
						(0.91)
$\Delta \bar{r}_t$						5.35***
						(1.19)
$\pi_t^{US,trailing\ 12m}$						-1.22*
						(0.67)
$\bar{\pi}_t{}^{trailing\ 12m}$						2.77***
						(0.99)
$\Delta GZ\ spread$						-0.03***
						(0.008)
$\Delta Treasury\ premium$						-0.0005*
						(0.0003)
Obs.	50	50	50	50	50	50
R^2	0.32	0.08	0.15	0.15	0.15	0.64
Shapley value (traditional)						51.3%
Shapley value (Δr_t^{US})						24.4%
Shapley value ($\Delta \bar{r}_t$)						16.2%
Shapley value ($\pi_t^{US,trailing\ 12m}$)						6.4%
Shapley value ($\bar{\pi}_t{}^{foreign,trailing\ 12m}$)						4.3%
Shapley value (financial / risk)						48.7%
Shapley value ($\Delta GZ\ spread$)						36.6%
Shapley value ($\Delta Treasury\ premium$)						12.1%

*Note: Column (1) to (5) are directly taken from LMNS. Column (6) is the Engel Wu model we produced. Standard errors in parentheses. * p<0.1, ** p<0.05, *** p<0.01.*

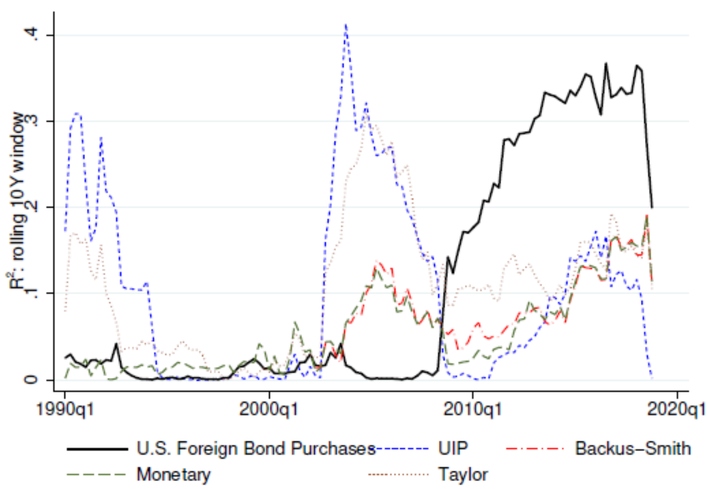
Column (6) shows that the regression has an R^2 of 64% at the quarterly frequency. All coefficients are statistically significant and have the expected signs. This indicates that the macroeconomic relationships we document are robust to modest changes in the dataset, frequency, and sample period. The Shapley value of the monetary variables (home and foreign interest rates, home and foreign inflation) and non-monetary variables (GZ spread and convenience yield) are 51.3% and 48.7% respectively, indicating the monetary variables are important and contribute to half of the explained variation (half of the R^2 of 64%).

2) Figure A.6 of LMNS

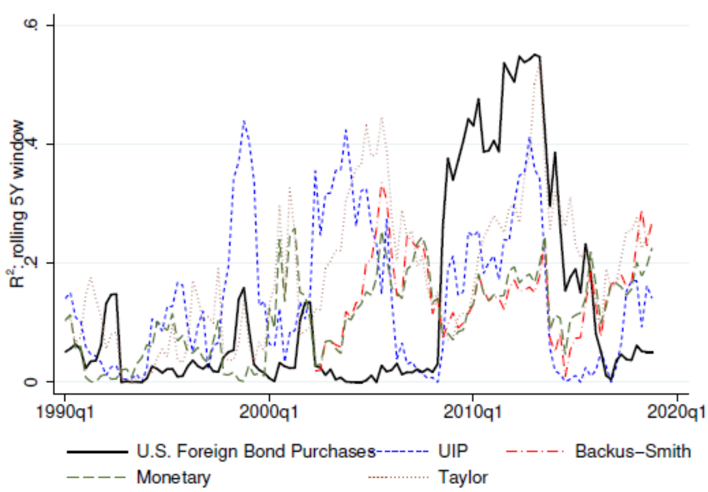
Using these in-sample regressions, LMNS report rolling-window estimates in their Figure A.6, which we reproduce below. We construct analogous rolling-window regressions using our baseline specification. Specifically, we report 5-year and 10-year rolling regressions, as in LMNS, and additionally include 20-year rolling-window results, as presented in the main text.

LMNS Figure A.6: In-sample Explanatory Power of Capital Flows and Other Fundamentals

(a) 10-Year Rolling Window



(b) 5-Year Rolling Window



Similar to Figure A.6 in LMNS, the figures below present rolling-window regressions, where each observation is indexed by the end date of the window on the x-axis. The black vertical line marks 1999Q1. We exclude the 2008–2009 period from the entire analysis to demonstrate that the results are not driven by the inclusion of the global financial crisis.

Across all three rolling windows, the average R^2 in the pre-1999 sample (red dashed lines) is lower than in the post-1999 sample (green dashed lines). For the 10-year and 20-year rolling windows, the R^2s are more stable, and we observe a further increase after 2009 (yellow dash lines). This suggests a strengthening relationship between exchange rates and macroeconomic variables.

Figure A12: in sample rolling window R^2 of Engel Wu model

Note: Blue lines are the rolling window R^2. Red dashed lines are the average R^2 value pre-1999. Green dashed lines are the average R^2 value post-1999. Yellow dashed lines are the average R^2 value post-2008.

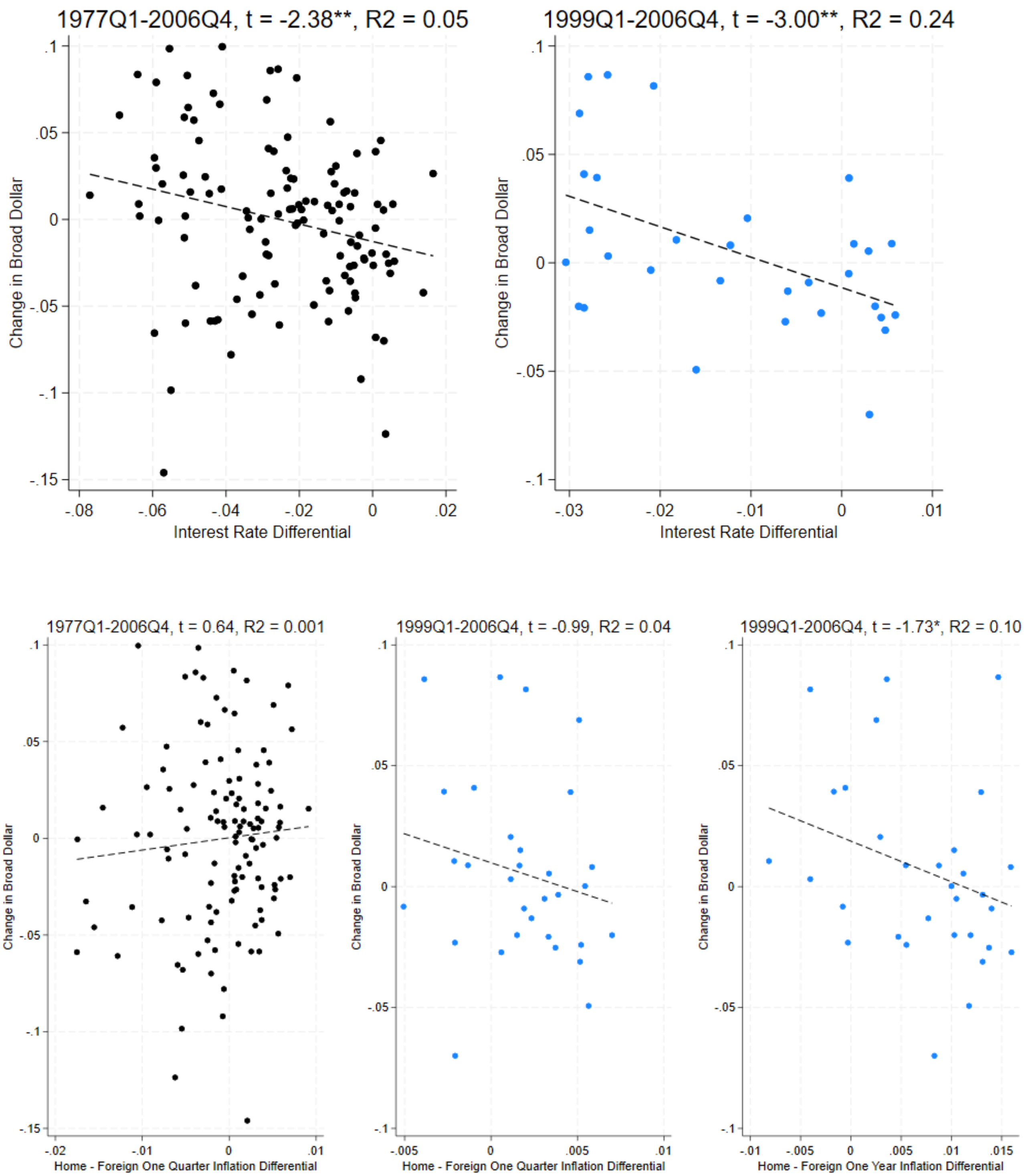
3) Figure A.1 of LMNS

In Figure A.1 of LMNS, the authors plot changes in the simple average exchange rate on the y-axis against either the interest rate differential or the inflation differential on the x-axis. We reproduce these results on the left-hand side of Figure 2 (black dots, interest rate differential in the upper panel, inflation differential in the lower panel), using a quarterly sample from 1977Q1 to 2006Q4.

As noted in the paper, the relationship between exchange rates and macroeconomic variables appears to shift around 1999. On the right-hand side of Figure 2, we plot the same relationships but restrict the sample to 1999Q1–2006Q4. The relationship between exchange rates and the interest rate differential becomes more negative, with an R^2 of 24%.

In the lower panels, we present the corresponding results for inflation differentials over the 1999Q1–2006Q4 period. The middle panel uses quarter-on-quarter inflation differentials, while the right panel uses year-on-year inflation differentials. In both cases, the relationship becomes more negative when focusing on the post-1999 sample.

Figure A13: reproduction of scatter plot of LMNS Figure A.1



Note: Left panels with black dots are directly taken from LMNS Figure A.1. Right panels (blue dots) are reproduced with sample restricted to post 1999.

4) Table A.5 of LMNS

Table A.5 of LMNS investigates the out-of-sample fit of the selected macroeconomic models, as in Meese Rogoff (1983). We reproduce these results below in Table A18, from row (a) Uncovered Interest Parity to row (d) the Taylor rule specification, using a 10-year rolling estimation window. Their reported results indicate no evidence that the selected macro models achieve a significantly lower RMSE than a random walk benchmark; all p-values exceed 0.10 (the row "Ratio" represents the RMSE of the model relative to that of the random walk benchmark.)

In row (e), we repeat the same out-of-sample fit exercise using our baseline regression specification. We exclude Japan and Switzerland when constructing the average exchange rate and foreign variables. This specification yields a significantly lower RMSE relative to the random walk benchmark, indicating that our baseline model delivers meaningful out-of-sample predictive power. The associated p-value falls below conventional significance thresholds, allowing us to reject the null hypothesis of equal predictive accuracy.

Table A18 Reproduction of Lilley et al Table A.5 and our baseline regression

		RMSE	Quarterly **All periods**	Quarterly **Without Crisis (exclude 2008Q1-2009Q2)**
			(1)	**(2)**
(a)	Uncovered interest parity	Ratio	1.04	1.05
		p-value	0.30	0.08
(b)	Backus Smith	Ratio	0.99	1.00
		p-value	0.81	0.85
(c)	Monetary	Ratio	0.99	1.01
		p-value	0.72	0.80
(d)	Taylor	Ratio	1.03	1.05
		p-value	0.69	0.39
(e)	Engel Wu specification	Ratio	0.84	0.96
		p-value	0.01**	0.03**

Note: Ratio represents the RMSE of the model relative to that of the random walk benchmark. p-value is Diebold-Mariano test for the performance of the model relative to the random walk model

Taken together, these results suggest that incorporating real interest rate differentials, inflation dynamics, and proxies for risk premia and convenience yields substantially improves exchange rate fits relative to the standard random walk benchmark. Furthermore, they provide evidence of a strong connection between exchange rates and macroeconomic fundamentals when focusing on the post-1999 sample and adopting a richer regression specification.