Presidential Cycles and Exchange Rates*

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Presidential Cycles and Exchange Rates

Abstract

This paper shows that US presidential cycles can predict dollar-based exchange rate returns. Armed with nearly 40 years of data and a large cross-section of currency pairs, we document an average US dollar appreciation during Democratic presidential terms and an average US dollar depreciation during Republican presidential mandates. The difference in these average exchange rate returns is larger than 5% per annum and is primarily linked to trade tariffs. In contrast, we find no relationship with cross-country interest rate differentials, inflation differentials, and pre-existing economic conditions. We relate these findings to trade policy within a model of exchange rate determination with constrained financiers.

Keywords: Presidential Cycles, Foreign Exchange, Currency Risk Premia, Tariff, Trade Policy Uncertainty.
JEL Classification: E44, D72, F13, F31, G15, G20, P16.

"The US election has the potential to be a significant market mover"

— Financial Times, September 28, 2020

1 Introduction

Exchange rates are notoriously difficult to forecast and there is limited empirical support for traditional models based on economic fundamentals. The forecasting power of these models is generally poorer than a simple random walk process (e.g., Meese and Rogoff, 1983; Engel and West, 2005). But exchange rates are affected by much more than just interest rates and inflation, their dynamics is extremely complex and usually puzzling (e.g., Engel, 2016; Hassan et al., 2023), and unexpected exchange rate shifts often happen around elections, referendum, and other political events. The connection between politics and foreign exchange markets, however, is not well understood and the empirical evidence remains scant being political factors not easily measurable. Not surprisingly, financial economists and practitioners are often caught off-guard when exchange rates are hit by major political events as *political information* may not be processed as efficiently as *economic information* (e.g., Roberts, 1990; Freeman et al., 2000).

In this paper, we study the relationship between exchange rates and US presidential cycles. On the one hand, the selection of the US president is a major political event that attracts massive global interest since the new president can reshape the foreign policies of the US, a country that is undoubtedly central to international trade and capital flows. On the other hand, US presidential elections are periodically scheduled and this regularity makes US political cycles uncontroversially easy to determine. Specifically, a presidential cycle starts when a political party gains victory at the presidential election and ends when the candidate of a different political party wins the White House. To preview our results based on nearly 40 years of data, we find that the US dollar is systematically stronger during Democratic presidential mandates than Republican presidencies relative to a large cross-section of developed and liquid emerging market currency pairs. On average, the US dollar appreciates by 4.31% per annum during Democratic presidential terms and depreciates by

1.25% when the US president is a Republican. The difference in average exchange rate returns (the return difference, henceforth) between Democratic and Republican presidential cycles is larger than 5% per annum, a figure that is both statistically significant and economically large.

A large body of the early literature examines the role of US presidential cycles for macroeconomic outlook and concludes that output growth is slower during Republican administrations whereas inflation rate is higher under Democratic presidencies (e.g., Alesina and Rosenthal, 1995; Alesina et al., 1997; Blinder and Watson, 2016). In contrast, only a few recent papers have studied the relationship between US presidential cycles and the performance of financial markets (e.g., Santa-Clara and Valkanov, 2003; Brogaard et al., 2020; Pástor and Veronesi, 2020). In particular, Santa-Clara and Valkanov (2003) is the first paper to document a higher excess return for US stock markets under Democratic than Republican presidencies, a stylized fact described as the 'Presidential Puzzle' due to the lack of plausible empirical explanations. Pástor and Veronesi (2020), moreover, attempt to rationalize this finding using a theoretical model that incorporates US tax policy and time-varying risk aversion. Brogaard et al. (2020), in addition, study the impact of global political uncertainty measured using US election cycles on global asset prices and uncover a strong negative empirical relationship. Liu and Shaliastovich (2017), finally, focus on US government policy approval arising from US presidential or congressional ratings and find a strong relationship with fluctuations in dollar exchange rates.

We also check empirically whether foreign political cycles can generate sizable exchange rate return differences. We generally find that foreign political cycles are statistically insignificant but the sign of the policy coefficients can be cross-sectionally inconsistent, potentially due to the irregularity and endogeneity of the election day. Therefore, it is difficult to reach a conclusion that the conventional bipartisan hypothesis applied to foreign countries can play any role in our analysis. Only US presidential cycles generate consistent and significant exchange rate return differences. These results are not confined to developed currencies but further extend to a range of liquid emerging market currencies, are not offset by the cross-country interest rate and inflation differentials, and are unrelated to traditional variables used to proxy for the US business cycle fluctuations such as the term spread, default spread, relative interest rate, and log dividend-price ratio (e.g., Santa-Clara and Valkanov, 2003).

Next, we investigate a possible explanation that rationalizes the causes of the observed exchange rate return difference. Trade policy is a natural candidate in the international context and the US president plays a special role compared to other presidents or prime ministers in foreign countries. There is some evidence that the US president can bypass congress to impose a tariff on imports, thus justifying our focus solely on the presidential cycle rather on congressional characteristics. We analyze the influence of trade restrictions on the dynamics of exchange rates and find that the degree of protectionism, not only measured by tariff levels but also by other forms of restrictions on the trades and payments, is important to understand the US presidential cycle for exchange rates.

Finally, we build and contribute to this recent literature by studying the relationship between US presidential cycles and exchange rate returns. We hypothesize that the US trade policy is implemented in significantly different directions between the two parties. Furthermore, we also consider the retaliatory trade restrictions imposed by foreign countries which results in a worldwide elevation of trade restrictions following the trade policy initiated by the US president. In the spirit of Gabaix and Maggiori (2015) we propose a three-period model of exchange rate determination in which two features are important for generating predictions aligned with the empirical findings. First, trade restrictions represented by the nonnegative tariff persists up to medium term, and second, the financiers have only constrained risk-bearing capacity. We find that a short-lived tariff shock can be absorbed by the financiers and thus is not sufficient to result in the US dollar depreciation. However, trade restrictions worldwide persisting up to the medium term can overload the limited risk-bearing capacity of the financiers and eventually lead to the US dollar depreciation. The model further predicts a US dollar appreciation when the ease of trade restrictions, described by the falling tariffs, converge to the frictionless equilibrium in the long run. These model implications match well our observations from the empirical data.

In short, our contribution can be summarized in three dimensions. First, we establish a connection between the US presidential cycle and exchange rates. We demonstrate that Democrats-Republicans presidential cycles, irrespective of foreign political cycles, contribute to an economically sizeable return difference for a large cross-section of currency pairs. Second, we rule out the possibility that the exchange rate return difference can be attributed to pre-election economic conditions. Finally, we propose trade policy as a plausible explanation, which we then verify in the data and further rationalize in a theoretical model. We show that trade centrality amplifies the exchange rate return difference.

Related Literature. The return difference is well known in the US stock market. Santa-Clara and Valkanov (2003) ruled out some potential explanations and documented this phenomenon as 'presidential puzzle'. They did not find significant relations between stock returns and congressional variables while similar findings were reported by Blinder and Watson (2016) over economic growth and congressional variables. The model in Pástor and Veronesi (2020) focused on the imposed tax policy under the Democratic and Republican presidents. They rationalized the return difference in the US stock market by providing an explanation of fiscal policy and time-varying risk premia. The influence of political uncertainty on the risk premia was well-explored with the US data. Pástor and Veronesi (2013) proposed a general equilibrium model to rationalize the price dynamics responding to political news.

Our paper is also related to the bipartisan models (e.g., Hibbs, 1977) and the political real business cycle (e.g., Nordhaus, 1975), in which Democrats prioritize growth over inflation and unemployment while Republicans favor the opposite. Alesina and Roubini (1992) investigated 18 OECD economies to document the long-run bipartisan differences in inflation and the temporary bipartisan differences in output and unemployment. On the contrary, Blinder and Watson (2016) documented the bipartisan difference in the US economic growth. Hence, it remains inconclusive whether the bipartisan hypothesis is an important factor to understand the international stock and currency market.

Lohmann and O'Halloran (1994) showed empirical evidence of lower tariff under a Democratic president. They also showed a similar finding under a unified government where the President is in the same party as the House and Senate majority. Recently, Fajgelbaum et al. (2019) and Fetzer and Schwarz (2020) investigated the economic losses due to the trade war raised by President Trump's administration via a specific dimension of the retaliation tariff enacted by the US trade partners. Our paper complements theirs by documenting that the tariff rate is an important factor to explain also the financial returns in a longer sample. On the other hand, Liu and Shaliastovich (2017) argued the relations between the policy approval and currency risk premium. They showed a higher rate of policy approval predicts higher economic growth and lower currency risk premium. Furthermore, Hassan et al. (2023) showed that the trade deals can increase trade substantially and reduce exchange rate's systematic risks. Our paper echoes their works by relating the trade policy and currency returns. We further propose the trade restrictions worldwide as a potential explanation for the return difference uncovered in the international currency markets.

Moreover, the influence of political uncertainty on the foreign exchange market was discussed in a few papers. Bachman (1992) offered an information-based explanation, showing that the forward exchange premium is mitigated after general elections in several industrial countries. With a focus on the exchange rate volatility, Lobo and Tufte (1998) analyzed the bipartisan effect in a sample of four countries. As for the study of international stock returns, our paper is related but different from the work by Brogaard et al. (2020). They studied the impact of political uncertainty on the international asset prices but the focus was to show the negative influences from the pre-election uncertainty rather than to understand the policy choice throughout the presidential terms. It is an extension of the single-country study reported in Kelly et al. (2016), which uses option data to verify the link between political uncertainty and risk premia. Our goal is beyond this type of event study which treats election as an exogenous shock. We aim to explain the political-economic fluctuations in the risk premia of the stock and foreign exchange markets.

The remaining of the paper is organized as follows. Section 2 describes our data sources. Section 3 presents the empirical analysis to document the spillovers of US presidential cycles to foreign exchange rates. Then, we formulate a model of exchange rate determination on trade policy uncertainty in Section 4 for the return difference before concluding in Section 5. A separate Internet Appendix provides additional robustness tests and supporting analysis.

2 Data and Preliminary Analysis

This section describes the main data employed in the empirical analysis and provides some preliminary results.

2.1 Data on Exchange Rates

Data on the daily spot and one-month forward exchange rates relative to the US dollar are sourced from Barclays Bank International and WM Reuters via Datastream. The empirical analysis employs monthly observations that we obtain by sampling end-of-month exchange rates between October 1983 and October 2020. We focus on a sample that includes the currencies of developed countries as well as the currencies of major emerging economies, i.e., Australia, Belgium, Brazil, Canada, Czech Republic, Denmark, Euro Area, France, Germany, Hungary, Italy, Japan, Mexico, Netherlands, New Zealand, Norway, Poland, Singapore, South Africa, South Korea, Sweden, Switzerland, Taiwan, Turkey, and the United Kingdom. After the introduction of the euro in January 1999, we drop Belgium, France, Germany, Italy, and the Netherlands from the sample. The sample starts with 9 currencies at the beginning of the sample in 1983 and ends with 20 currencies at the end of the sample in 2020.

Exchange rates are expressed in units of US dollar per unit of foreign currency so that an increase in the exchange rate indicates an appreciation of the foreign currency or equivalently a depreciation of the US dollar. We define spot and forward exchange rates at time t in logs as s_t and f_t, respectively. Monthly exchange rate returns from buying a unit of foreign currency in at time t while reversing the position at time $t+1$, both in the spot market, are denoted as $\Delta s_{t+1} = s_{t+1} - s_t$. Similarly, monthly excess returns from buying a unit of foreign currency in the forward market at time t and then selling it in the spot market at time $t+1$ are computed as $rx_{t+1} = s_{t+1} - f_t$. We also construct real exchange rate returns between months t and $t+1$ as $\Delta q_{t+1} = \Delta s_{t+1} + \pi^*_{t+1} - \pi_{t+1}$, where π_{t+1} and π^*_{t+1} are the inflation rates for the US and the foreign country, respectively, between months t and $t+1$. We collect monthly observations on year-on-year inflation rates from Datastream and suitable scale them to proxy for π_{t+1} and π^*_{t+1}.

2.2 Data on Political Variables

Data on the US presidential cycles are hand collected and span, respectively, six Republican presidential terms and four Democratic presidential terms. The latter includes the presidencies of Bill Clinton and Barack Obama whereas the former comprises the presidencies of Ronald Reagan, George H.W. Bush, Gorge W. Bush, and Donald Trump. Each cycle starts in November when the US presidential election takes place and ends four years after in October. Based on these data, we define DP_t, a monthly time series of a Democratic dummy variable, that takes on the value of one during a Democratic presidential cycle and zero under a Republican presidential cycle. For example, under the presidential terms of Barack Obama, DP_t is set equal to one between November 2008 and October 2016. Overall, our sample combines 192 months of Democratic presidential terms (or 43.1% of all months) and 253 months of Republican presidential terms for a total of 445 months.

We also collect data on the election cycles of other major economies, which we use to define control dummy variables. Specifically, we source the information from the latest Database of Political Institutions (DPI) by Scartascini et al. (2020) with the adjustments of our own hand-collect data such as election dates. We focus on the members of the G7 countries, i.e., Canada, France, Germany, Italy, Japan, and the United Kingdom, and broadly categorize the winning party (or ruling coalition) in each country either as a center-left or center-right political party. In particular, we classify the Liberal Party as center-left and the Conservative Party as center-right in Canada; the Social Democratic Party as center-left and the CDU/CSU as center-right in Germany; the Socialist Party and En Marche! as center-left, and the Republicans, Rally for the Republic, and UMP as center-right in France; the Labour Party as center-left and the Tory Party as center-right in the United Kingdom; the Pentapartito, Olive Tree, Union, Democratic Party, and the coalition Democratic Party, Five Star Movement, and Free and Equal as center-left, and the Pole of Freedom, House of Freedom, People of Freedom, and the coalition Five Star Movement and Lega as center-right in Italy; the Democratic as center-left and the Liberal Democratic as center-right in Japan.

For each country, we then define CL_t, a monthly time series of a control dummy, that is equal to one when the winning party (or the ruling coalition) is leaning towards the center-left and zero otherwise. Each cycle starts when we observe new elections or a swing in the ruling coalition. Note that if the election takes place in the second half of the month, we begin the cycle only in the following month. As an example, take the last two general election cycles in Italy. We assign a value of zero starting from March 2013, which is the month following the general election that was held on the 24^{th} (second half) of February in 2013. On the other hand, we switch our dummy to one starting from the same month, March 2018, since the election took place on the 4^{th} (first half) of March in 2018.

Table 1 about here

In Figure 1, we summarize the percentage number of times that a dummy variable is equal to one or zero. In Canada, the dummy is equal to one in 50.3% of all months, thus indicating an equal split between center-left and center-right. France and Italy, moreover, are leaning towards the center-left as the dummy equals one in 54.2% and 64.0% of all months, respectively. On the contrary, Germany, Japan, and the UK are drifting towards the center-right since the dummy amounts to one in 18.9%, 15.1%, and 35.1% of all months, respectively.

2.3 Data on Trade Variables

We sample trade data on the aggregate imports and exports of goods and services[1] from the International Monetary Fund's Direction of Trade Statistics (DOTS) as well as data on GDP via the Datastream platform. Data are expressed in US dollar and range at quarterly frequency.

[1]In the earlier version of this paper, we analyzed the data of Custom and Duties from the World Bank. There were much more missing data both on the time-series and cross-sectional dimensions. Specifically, 12 out of 25 countries do not report the custom and duties. In addition, there were many zeros and even negative values making it difficult to interpret and impossible for the log transformation, which is usually useful for the execution of cross-country comparison. Hence, we search for alternative data that can reflect the trade barriers in a broader sense in order to avoid the above-mentioned issue.

2.4 Data on Trade Restrictions

We collect the Measure of Aggregate Trade Restrictions (MATR) from the IMF, a newly proposed measure of trade policy by Estefania-Flores et al. (2022).[2] The MATR is recorded at annual frequency, available for 157 countries and 70 years from 1949 to 2019. It is constructed based on the narrative analysis of policy using data from the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions. Although the panel data remains unbalanced but it is much more complete than some existing measures, such as the trade tariffs. Note that the MATR codes the existence of trade restrictions rather than their intensity, similar to the capital control index proposed by Chinn and Ito (2008).

It is also important to employ a measure broader than the tariffs in our study because of the following two reasons: first, as most countries are member of World Trade Organization, their commitment to keep low custom duty rates results in the tariff data lack of variation in time series and cross section. Note that many countries report 0 in the custom duties in the recent years; Second, the tariff is merely one of the many measures to impose trade restrictions. Estefania-Flores et al. (2022) has documented the large and negative correlation between the MATR and the economic output. We further study its impact in the foreign exchange rate markets.

We choose to work with the MATR index that includes the tariff barriers since the index with non-tariff barriers tracks closely to the former. In our sample, only one currency TWD do not have the MATR observations. Overall, the MATR should be understood as a score of trade restrictions, ranging from 2 to 19 in our sample. Because the scale varies across countries, we take log on the MATR and focus on the difference with respect to the previous year to measure the change of trade policy in each country.

[2] The literature has investigated the macroeconomic impact of trade policy by using the tariff data in the bipartisan context (Gardner and Kimbrough, 1989; Lohmann and O'Halloran, 1994). Nevertheless, the trade policy has evolved to a wide variety of measures other than tariffs during the past decades. More recent literature uses more trade "barrier" or "restriction" in the study of trade policy.

2.5 Data on US Macroeconomic Variables and Fiscal Policy

We also collect data on a variety of US macroeconomic variables that we use to proxy for business cycle fluctuations akin to Santa-Clara and Valkanov (2003). This set of variables included the log dividend-price ratio LDP_t, the term spread TSP_t between the ten-year Treasury constant maturity rate and the three-month Treasury bill rate, the default spread DFS_t between yields of BAA-rated and AAA-rated corporate bonds, and the relative interest rate RR_t computed as three-month Treasury bill rate in deviation of its one-year moving average. For all these data, we obtain end-of-month data by All these data are monthly and The dividend-price ratio is available from Robert Shiller's website whereas the other data are obtained from the FRED database maintained by the Federal Reserve Bank of St. Louis. Last but not least, to measure the US fiscal policy, which is proposed by Pástor and Veronesi (2020) as a major explanatory variable to rationalize the presidential puzzle in the US stock market, we sourced the federal tax revenue and the US GDP at quarterly frequency also from the FRED database.

3 Main Findings: Democrats versus Republicans

This section shows that exchange rate returns comove with the US presidential cycles. Using a large cross-section of currency pairs, we document that the US dollar tends to appreciate during Democratic presidential terms and depreciate under Republican presidential terms. The difference in dollar-based exchange rate returns between Democratic and Republican presidencies is statistically significant, can be attributed neither to interest rate differential nor to the inflation differential, and is not driven by fluctuations in US business cycle variables. IN contrast, this difference can rationalized using trade policies and tariff uncertainty.

3.1 Exchange Rate Return Performance

We establish our findings by first presenting summary statistics of country-level monthly exchange rate returns. Table 1 reports the means and standard deviations in percentage per annum for the full sample that ranges between October 1983 and October 2020 as well as Democratic and Republic presidential terms. The former subsample is denoted as DP whereas the latter is referred to as RP.

TABLE 1 ABOUT HERE

The first two columns of Table 1 refer to the full sample, which includes 445 months. Recall that exchange rates are defined as units of US dollar per unit of foreign currency and a negative return indicates an appreciation of the dollar. Out of 25 currency pairs, 14 currency pairs have experienced depreciation and 11 currency pairs have gone through an appreciation against the dollar. With a few exceptions, mainly concentrated around emerging market economies like Brazil, Mexico, Turkey, and South Africa, there is no clear pattern on whether the US dollar has on average appreciated or depreciated against foreign currency pairs during our sample. We further add means and standard deviations of an equally-weighted basket (EWR) and a GDP-weighted basket (VWR) and get to the same conclusion. The EWR basket displays an average exchange rate return that is slightly negative (-1.15% per annum) whereas the VWR basket shows an average exchange rate return that is indistinguishable from zero (-0.05% per annum). The exchange rate volatility, moreover, evolves around 12% for individual currency pairs and is slightly above 8% for the currency baskets.

The next two columns of Table 1, under the heading of DP, report the summary statistics for Democratic presidential terms, a subsample that includes 192 months. With the single exception of the Japanese yen, the US dollar has on average appreciated against all other currency pairs during Democratic presidential terms. This stylized fact is further confirmed when currency pairs are grouped together. The EWR basket exhibits an average US dollar appreciation of 4.31% per annum, which is economically sizeable and three percentage points larger than the corresponding figure reported for the full sample. We uncover similar results for the VWR basket, i.e., an average US dollar appreciation of 3.12% per annum that is three percentage points larger than the corresponding full-sample statistic. The columns under the heading of RP, in contrast, denote the Republican presidential terms, a subsample that is slightly larger and comprises 253 months. We find that, under Republican presidents, the US dollar has on average depreciated against 19 out of 25 currency pairs in our sample. The cross-country baskets, moreover, point towards the same conclusion since the EWR and VWR basket display an average US dollar depreciation of 1.25% and 2.29% per annum, respectively. These results, taken together, suggest that the US dollar on average appreciates under Democratic presidents and depreciates under Republican presidents.

FIGURE 2 ABOUT HERE

In the last two columns of Table 1, we show the mean and standard deviation differences between Democratic and Republican presidential terms. Except for the Brazilian real, the mean difference is always negative and evolves around -5.54% per annum for the developed currency pairs (i.e., the first 15 of the list) and -6.25% per annum for emerging market currency pairs (i.e., the last 10 of the list). These findings can be further visualized in the bar chart reported in Figure 2, which also shows that there is more cross-country variation for emerging market currencies than developed currencies. The mean differences for currency baskets, moreover, are virtually identical since EWR displays a mean difference of -5.56% per annum while VWR exhibits a mean difference of -5.42% per annum. Finally, while the exchange rate volatility is on average lower under Democratic presidential terms than under Republican presidential terms, its difference is economically small and slightly larger than 0.50% in absolute terms.

Overall, this first set of results documents a striking regularity that characterizes dollar-based exchange rate returns: the US dollar on average appreciates during Democratic presidential terms and depreciates during Republican presidential terms. We thus complement the work of Santa-Clara and Valkanov (2003) and Pástor and Veronesi (2020), who show that the average US stock market excess return is higher under Democratic than Republican presidencies.

3.2 The Role of Interest Rates

The findings reported in the previous section beg the question of whether our results are driven and, to some extent, offset by cross-country interest rate differentials. We run two different exercises to verify this legitimate concern. In the first exercise, we first replace the country-level exchange rate returns with country-level currency excess returns and then compute summary statistics for the full sample as well as Democratic and Republican presidential cycles.

Table 2 about here

We present our results in Table 2 and uncover no substantial difference relative to our core results. In particular, the mean differences between Democratic and Republican presidential terms reveal that our findings remain robust to the inclusion of the interest rate differential in 22 out of 25 currency pairs. Except for three emerging market currencies, i.e., the Brazilian real, Mexican peso, and South Korean won, the mean difference is always negative and moves around -6.15% per annum for the developed currency pairs and -3.24% per annum for emerging market currency pairs. These results suggest that for developed currencies there is virtually no difference on average between exchange rate returns and currency excess returns. For emerging market currencies, however, local interest rates are slightly higher on average under Democratic presidents than Republican presidents. A plausible explanation is that Central Banks in emerging market countries are likely to respond to local currency depreciation by raising short-term local interest rates. This consideration, however, mildly affects our overall results as the EWR and VWR baskets continue to exhibit an economically large mean difference of about -4.37% and -5.05% per annum, respectively.

Figure 4 about here

In the second exercise, we calculate the exchange rate returns of a pseudo trading strategy that naïvely buys the US dollar while shorting an equally-weighted basket of foreign currencies under a Democratic White House and sells the US dollar while investing in an equally-weighted basket of foreign currencies under a Republican White House. We then compare the exchange rate returns of this pseudo strategy, labeled as the 'dollar cycle', with the exchange rate returns of the 'dollar carry' of Lustig et al. (2014) and Verdelhan (2018). The latter is an investment strategy that exploits the time-series variation in the average US interest rate difference relative to the foreign countries. It takes a long position in an equally-weighted basket of foreign currencies while selling the US dollar whenever the average foreign short-term interest rate is above the short-term US interest rate and sells an equally-weighted basket of foreign currencies while going long the dollar whenever the short-term US interest rate is higher than the average foreign short-term interest rate. We plot the cumulative exchange rate returns of these two strategies in Figure 4 and their time-series behaviors look remarkably different: the 'dollar cycle' yields an average exchange rate return of 2.6% per annum whereas the 'dollar carry' produces an average exchange rate return of −0.4% per annum. These figures coupled with a return correlation that is as low as 1% suggest that the interest rate differential is unlikely to be a primary driver of our core results.[3]

The 'dollar cycle' has predicted the average exchange rate return considerably well between its inception and late 2008. It has then struggled during the global financial crisis that followed the Lehman Brothers collapse before turning on a positive drift again between mid-2011 and early 2018. At that time, the Tax Reform passed by the Trump administration introduced an incentive for US firms to repatriate their offshore cash holdings and likely acted as a major source of demand for the US dollar.[4] The more recent COVID-19 outbreak beginning in the late March of 2020 and the associated flight-to-safety behavior of global investors can further explain the recent US dollar appreciation and the resulting negative

[3]The comparison is based on exchange rate returns solely to verify whether our core results are driven by the interest rate differential. In terms of profitability, the strategies are largely comparable as the 'dollar cycle' generates an average currency excess return of 2.4% per annum whereas the 'dollar carry' delivers an average currency excess return of 2.1% per annum. Ranking these strategies in terms of profitability, however, is beyond the scope of this exercise.

[4]Since there were a few tax-related bills passed around the same period, we clarify by referring the Tax Reform to the *Tax Cuts and Jobs Act*, passed with no support from the Democratic Party and signed by the President in December 2017. This bill affects the US international businesses as well as the US citizens living and working abroad. Furthermore, the US dollar also rose after President Bush offered a tax holiday on repatriated earnings in 2004 with the *Homeland Investment Act*.

performance of our pseudo trading strategy. The 'dollar carry trade', in contrast, has predicted the average exchange rate return reasonably well at the beginning of the sample before weakening its predictive power since the early '90s.

FIGURE 3 ABOUT HERE

We also construct both strategies using GDP-weighted baskets of foreign currencies but results remain qualitatively similar albeit with a marginally higher sample return correlation of about 7%. Cumulative exchange rate returns are displayed in Figure 3. Overall, it seems that interest rates are unlikely to be the main determinant of our results.

3.3 The Role of Inflation Rates

We also check whether our findings can be attributed to cross-country inflation rate differentials. To shed light on this question, we carry out two different exercises similar to those presented in the previous section. In the first exercise, we first replace the country-level nominal exchange rate returns with country-level real exchange rate returns and then present summary statistics for the full sample as well as Democratic and Republican presidential mandates.

TABLE 3 ABOUT HERE

We show our results in Table 3 and find that the mean difference between Democratic and Republican presidential terms remain negative for 21 out 25 currency pairs. Excluding four emerging market currencies, i.e., the Brazilian real, Mexican peso, South Korean won, and Turkish lira, the mean differences between Democratic and Republican presidential terms are always negative. On average, it is about −5.75% per annum for the group of developed currencies and −4% per annum for both developed and emerging market currency pairs. Also, the EWR and VWR baskets display a mean difference of about −3.62% and −4.56%

per annum, respectively. These results thus suggest an average US dollar appreciation in real terms under a Democratic White House.

We also compare the exchange rate returns of our 'dollar cycle' pseudo strategy with the exchange rate returns of a 'dollar value' strategy that exploits the time-series variation in the average inflation rate difference between the US and the foreign countries in the spirit of Asness et al. (2013). It takes a long position in an equally-weighted basket of foreign currencies while selling the US dollar whenever the US inflation rate is above the average foreign inflation rate and sells an equally-weighted basket of foreign currencies while investing the dollar whenever the average foreign inflation rate is higher than the US inflation rate. We plot the cumulative exchange rate returns of the 'dollar value' strategy in Figure 4. This strategy delivers an average exchange rate return of -0.7% per annum and displays a return correlation of -12% with the 'dollar cycle' strategy. To sum up, the inflation differential is unlikely to fully offset the presidential cycle that characterizes dollar-based exchange rate returns.

3.4 Testing for the Presidential Cycle

We now carry a statistical assessment of the relationship between exchange rate returns and the US presidential cycle akin to Santa-Clara and Valkanov (2003), Specifically, we run regressions based on the following specification

$$\Delta s_{i,t+1} = \alpha + \beta\, DP_t + \varepsilon_t, \tag{1}$$

where $\Delta s_{i,t+1}$ is the exchange rate return for the currency i relative to the US dollar between months t and $t+1$, and DP_t is a presidential dummy variable that takes on the value of one (zero) during a Democratic (Republican) presidential terms assumed to be known at the start of the presidential cycle. We run both pooled and panel regressions with time-invariant currency fixed-effects. Under the null hypothesis that the presidential cycle does not affect exchange rate returns, we should obtain that $\beta = 0$. Differently, β will measure the mean exchange rate return difference between Democratic and Republican presidential mandates. Put differently, while α quantifies the average exchange rate return under a Republican White

House, the sum of α and β delivers the average exchange rate return under a Democratic presidential term.

TABLE 4 ABOUT HERE

We report estimates of α and β obtained via least-squares in Table 4 with standard errors clustered by currency and time (calendar date) dimension in parentheses. Panel A presents pooled regression estimates and documents a positive but statistically insignificant estimate of α (≈ 1.65 with a clustered standard error of 1.87) coupled with a negative and statistically significant estimate of β (≈ -5.78 with a standard error of 2.63). These estimates, given our definition of exchange rates, imply a statistically significant yet economically large appreciation of the US dollar ($\approx 4.13\%$ per annum) under Democratic presidential cycles and a statistically insignificant yet economically small depreciation of the US dollar ($\approx 1.65\%$ per annum) under Republican presidential terms. In Panel B, moreover, we absorb time-invariant unobserved currency characteristics but results are equivalent. In particular, the estimate of α (≈ 1.40 with a standard error of 1.75) is positive and statistically insignificant whereas the estimate of β (≈ -5.29 with a standard error of 2.54) is negative and statistically significant. Taken together, these estimates signify a statistically significant yet economically large appreciation of the US dollar ($\approx 3.90\%$ per annum) under Democratic presidencies and a statistically insignificant yet economically small depreciation of the US dollar ($\approx 1.40\%$ per annum) under Republican presidencies.

We also check whether our estimates are driven by the inclusion of a particular currency pair in our analysis. To this end, we sequentially remove one currency pair at a time before re-estimating the regressions implied by Equation (1). These estimates are reported in Table 4 but the results remain qualitatively similar, i.e., all estimates of α are positive but statistically insignificant whereas all estimates of β are negative and statistically significant regardless of whether we employ pooled or panel regression methods. In terms of economic value, on average, the US dollar appreciates the most during Democratic presidential terms when we drop the Japanese yen and the least when we exclude the Turkish lira. Our pooled (panel) regression estimates imply an average US dollar appreciation of about 4.27% (4.03%) per annum in the former case and an average US dollar appreciation of about 3.04% (2.94%)

per annum in the latter case. To sum up, we find that the relationship between exchange rate returns and the US presidential cycle is not only economically important but also statistically significant.

3.5 Controlling for Local Political Cycles

In the previous section, we have established the existence of a statistically significant difference in exchange rate returns between Democratic and Republican presidencies. We now investigate whether our results are correlated with local political cycles. For this exercise, we take the election cycles of the G7 countries into account and augment Equation (1) as follows

$$\Delta s_{i,t+1} = \alpha + \beta\, DP_t + \gamma\, CL_t + \varepsilon_t, \tag{2}$$

where CL_t is the control dummy variable that equals one when the winning party (or coalition) in the foreign country is leaning towards the center-left political spectrum and zero when the winning party (or coalition) has a center-right political agenda. When the prime minister does not come from the winning party, we use the prime minister's party to define our CL_t variable. The control dummy is defined for Canada, France, Germany, Italy, Japan, and the UK. A critical aspect of this exercise is that general elections in countries like Canada, Italy, Japan, and the UK may take place at irregular intervals and be endogenously driven by opportunistic political behavior (e.g., Goto et al., 2020). Additionally, for the elections taking place in the second half of the month, we update the control dummy only in the following month since the influence of election results is limited in the current month's exchange rate returns.

TABLE 5 ABOUT HERE

Table 5 presents pooled regression estimates of α, β, and γ with standard errors (in parentheses) by currency and time dimension. We find that our core results are not affected local political cycles and estimates of β are in line with those reported in Table 4. For example, when CL_t captures the political cycle in Germany, the estimate of α ($\approx$ 1.55 with a standard error of 2.08) and γ (≈ 0.38 with a standard error of 2.60) are both positive but statistically insignificant whereas the estimate of β is negative and statistically insignificant (≈ -5.76 with a standard error of 2.60). In Table A.1 in the Internet Appendix, we report panel regression estimates with time-invariant currency fixed effects but find no qualitative difference to our results. To sum up, adding control dummy variables that summarize local political cycles has a negligible impact on the correlation between exchange rate returns and the US presidential cycle.

3.6 The Role of Business Cycle Fluctuations

Political variables are often associated with business cycle fluctuations (e.g., Alesina et al., 1997; Drazen, 2000) and our findings may simply capture comovements between exchange rate returns and variations in the economic activity. If this is the case, the statistical significance recorded in the previous section should then weaken when variables that proxy for business cycle fluctuations in the US are taken into account. To test this hypothesis, we follow Santa-Clara and Valkanov (2003) and run predictive regressions based on the following specification

$$\Delta s_{i,t+1} = \alpha + \beta\, DP_t + \gamma'\, X_t + \varepsilon_t, \qquad (3)$$

where X_t denotes a set of predetermined macroeconomic variables, generally associated with the US business cycle, such as the term spread TSP_t, the default spread DSP_t, the relative interest rate RR_t, and the log dividend-price ratio LDP_t. If the political dummy variable only reflects information stemming from business cycle fluctuations, we should then observe a statistically insignificant and economically small estimate of β.

Table 6 about here

We report pooled regression estimates of α, β, and γ with standard errors (in parentheses) clustered by currency and time dimension in Table 6. In these regressions, all control variables are demeaned so that the coefficient estimates associated with DP_t are directly comparable with those reported in Table 4. Panel A presents different specifications based on control variables lagged by one month. The magnitude and the statistical significance of the β estimates, however, remain very similar to those without control variables, suggesting that the presidential dummy variable has an explanatory power for expected exchange rate returns that is largely orthogonal to proxies for US business cycle fluctuations. For example, specification (5) pulls all control variables together and produces a negative and statistically significant estimate of β (≈ -6.24 with a standard error of 2.80) that implies an average US dollar appreciation of 4.12% per annum under a Democratic president.

In Panels B through D, we verify the robustness of our results by increasing the lag of the control variables between three months and one year relative to the exchange rate returns. Overall, no significant difference is detected in our results. In Panel D, for example, we lag the control variable in X_t by one year. Specification (5) then yields a negative and statistically significant estimate of β (≈ -6.61 with a standard error of 2.75) that translates into an average US dollar appreciation of 4.65% per annum under a Democratic president. To conclude, similar to the evidence reported in Santa-Clara and Valkanov (2003), the results in Table 6 indicate that the correlation between exchange rate returns and political variables cannot be attributed to an indirect relation between business cycle fluctuations and presidential mandates.

3.7 The Role of Global Trade Policy

The foreign exchange markets are naturally linked to global trade, whose transactions are often invoiced in US dollars regardless of the countries involved in the trade (e.g., Gopinath et al., 2020). Studying the role of trade policies may then be important to rationalize the exchange rate return difference between Democratic and Republican presidential terms. Intuitively, trade policies that favors international trade would be associated with an increase in the demand for US dollars whereas trade policies that are more protectionist would go hand in hand with a decline in the demand for US dollars.

Tables 7 about here

The existing literature on trade and partisanship in political science acknowledges that politicians often take a stance on trade policy to please their constituents and win their election (e.g., Milner and Judkins, 2004). For example, Epstein and O'Halloran (1996) explore the US trade policy between 1877 and 1934 and report that Republicans (Democrats) enacted higher (lower) tariffs, even after controlling for economic factors. Irwin (2019), moreover, reaches similar conclusions by showing that US politics between 1861 and 1932 was dominated by Republicans who introduced higher tariffs to restrict imports. In contrast, Democrats reduced tariffs both in 1984 and 1913. From 1933 to 1993, in contrast, US politics was dominated by Democrats who favored trade agreements and lower tariffs. We uncover a similar pattern in our sample as shown by Table 7, where we list manually collected information about trade deals and trade disputes between 1983 and 2020. We find that the number of trade disputes under Republican Presidents is nearly double than under Democratic Presidents (19 under Republicans and 10 under Democrats). The number of trade deals, in contrast, is largely comparable (7 under Republicans and 5 under Democrats).

Motivated by this literature, we look for a broader measure which can capture the change of trade policy beyond the tariffs. With the MATR, we investigate the role of trade restrictions for exchange rate returns between Democratic and Republican presidential cycles by running panel regressions based on the following specification

$$\Delta s_{i,t+1} = \alpha_i + \beta_1 DP_t + \beta_2 MATR_{i,t} + \beta_3 \textit{Change of trade frictions}_{i,t} + \gamma' X_t + \varepsilon_{t+1}, \qquad (4)$$

where $MATR_{i,t}$ denotes the trade restrictions which is the log difference of the MATR index collected from the IMF expressed in percentage, *Change of trade frictions* is estimated from the detrended series of log imports and exports, X_t is a set control variables of alternative policies such as fiscal and monetary policies, and α_i refers to time-invariant currency fixed effects. While β_2 captures the correlation between trade restrictions and exchange rate returns under Republican presidents, the β_3 quantifies the general impact of the change of trade frictions on the exchange rates. We source annual data on the MATR from the IMF database and take difference after a forward-filled log transformation.[5] The set of control variables includes the US Fiscal Policy (a proxy calculated from the US federal tax revenue),

[5] A robustness test using the annual frequency data is conducted and the results are quantitatively similar.

and Monetary Policy (a proxy of the cross-country change on monetary policy).

Table 8 about here

We report the least-square estimates of our panel regressions in Table 8, while clustering standard errors (reported in parentheses) at the country and time (calendar year) dimension. In specification (1), we find a positive and statistically significant estimate of β_2 (≈ 0.526 with a standard error of 0.201), implying a US dollar depreciation in the growing of trade restrictions worldwide. Also, the estimate of β_1 associated with the presidential dummy becomes weakly significant, thus suggesting that part of its explanatory power is indeed captured by the role of trade policies. In specifications (2)–(3), we control for the changes of trade frictions by using the detrended imports and exports data from both the US and foreign countries. Note that the β_1 becomes insignificant while β_2 remains similar magnitude that is also positive and significant (≈ 0.498 and 0.614 with a standard error of 0.192 and 0.174, respectively), highlighting that the changes of trade frictions play an instrumental role in explaining the presidential puzzles in the exchange rate markets. Next, in specification (4)–(5), we examine the alternative explanations, including fiscal and monetary policies. Different from the findings in the US stock market in Pástor and Veronesi (2020), we do not find significant results in the US fiscal policy nor the cross-country difference of monetary policy. In the meanwhile, both β_2 and β_3 remain quantitatively similar to the previous three specifications.

Finally, in specification (6), we introduce *net import* as an alternative measure for the change in trade frictions of both the US and foreign countries. There are three main reasons of considering this alternative measure. First, the detrended trade variables may contain the business cycle fluctuations which have real-exchange-rate implications at least in a few countries studied by Bown and Crowley (2013). Taking the difference between import and export can eliminate the commonality due to the business cycle. Second, to keep the model tractable in the next section, we follow Gabaix and Maggiori (2015) to standardize the trade variables, condensing all the information down to one variable, the US net import. Hence, the choice of net import rather than net exports would also help us to validate the model predictions. Third, note that specifications (2)–(4) show a significant constant term, where the potential risk of collinearity among the variables of trade frictions calls for the necessity to test with an alternative variable. Although the US net import is found to be insignificant, the net import in foreign countries remain significant and negative (≈ -28.496 with a standard error of 11.487). In addition, the discovery of the insignificant constant alleviates the collinearity concern. The signs of both net imports in specification (6) are aligned with those in the previous specifications. More interestingly, the result is consistent with the model predictions that a Dollar appreciation is associated with a mitigation on the global trade frictions, approximated by a decreasing US net import ($\beta_3^{US} > 0$) or an increasing net import of foreign countries ($\beta_3^{Foreign} < 0$). Overall, the trade policy measures including both the aggregate and changes of trade restrictions appear to play an important role in explaining the exchange rate return difference between Democratic and Republican presidential terms.

4 Theoretical Framework

In this section, we rationalize our empirical findings by presenting a model of exchange rate determination with imperfect international trade in the spirit of Gabaix and Maggiori (2015).

4.1 Model

Consider a discrete-time model that lasts for three periods $t = 0, 1, 2$ in which the period 2 captures the long-run steady state while the period 1 describes a medium-term equilibrium. The economy consists of two countries, each populated by a continuum of households who produce and trade goods in an international market for goods and invest with financiers in risk-free bonds in their domestic currencies. There is a unit mass of global financiers who intermediate the capital flows resulting from households' decisions and absorb the currency imbalances at a certain level of risk premium. Without loss of generality, we refer to the domestic country as the US and its currency as the US dollar while the foreign country as Japan whose currency is the Japanese yen.

The innovation of our model is an imperfect trade market with frictions at global level. We consider a scenario where both the US and Japanese governments introduce trade restrictions on the imported goods, which results in lower income by their exporting producers. Thus, we model the *worldwide* trade restrictions by postulating that the policy actions and reactions of imposing trade restrictions tend to be positive correlated across countries. Eventually, the trade policy will affect the net fore asset balances and consequently the equilibrium exchange rates.

Financial intermediation is also imperfect because financiers are assumed to have limited capacity to absorb currency risk which leads to a downward-sloping demand curve for risk-taking. The equilibrium is achieved by a relative price (i.e., the exchange rate) in an international financial market. Essentially, the adjustment of the exchange rate clears the demand and supply of capital denominated in both currencies and, thus the exchange rate is determined in an imperfect capital market. In the following, we describe each of the model's players, their optimization problems, and analyze the resulting equilibrium.

4.1.1 Households

We assume that the maximization problem of households is similar across the countries. Hence, we only explain the details of the domestic households for the sake of brevity. The households need to solve an intertemporal consumption problem under the assumption of a logarithm utility function. In each period, the consumption is represented by a bundle of three elements: nontradable goods, domestic tradable goods, and foreign tradable goods. We adapt the Cobb-Douglas function to aggregate the three different goods. Note that the nontradable goods serve as the numéraire in each country so its price equals 1 in domestic currency.

The concept of *risk-free* security refers to a financial asset paying one unit of nontradable goods in all states of the world. The discount factor in each country is simply the reciprocal of the return on the domestic bonds. Selecting the consumption allocation between nontradable goods, and domestic and foreign tradable goods, the households maximize their utility subject to the market-clearing conditions for all goods which equate the total values of the production with consumption. We assume the production of both tradable and nontradable goods as exogenous.

We further assume that there is an absence of incentive for Japan to initiate the trade war against the US.[6] Nevertheless, Japan would retaliate for the rise imposed by the US. To put it explicitly into our model, we use a single variable τ_t^g to represent trade restriction worldwide, which is assumed to be a function of the US trade policy. Consequently, the Japanese government would also introduce the restrictions $\tau_t^g > 0$ for two periods consecutively $t = \{0, 1\}$, where the trade restrictions measures are expressed in a form of tariff as a percentage of the exporting goods value for the US household. Note that in the long run, the tariff will converge to zero, $\tau_T^g = 0$ for $T = 2$, implying that the trade restrictions will be removed eventually.

The first-order condition relevant for the tradable goods pins down the pre-tax value of the US exports as $\lambda_t p^*_{H,t} C^*_{H,t} = \xi_t/(1+\tau_t^g)$, where $C^*_{H,t}$ is the Japanese consumption of US goods and $p^*_{H,t}$ is its price. To keep the model most tractable, we set the shadow price of total production $\lambda_t = 1$, neutralizing the intertemporal variation in household marginal utility, which is not at the core of this paper. As a result, the dollar value of the after-tax imports boils down to $p^*_{H,t} C^*_{H,t} = \xi_t/(1+\tau_t^g)$[7].

The exchange rate e_t is defined as the number of dollar per unit of yen. Consequently, an increase in e_t implies a dollar depreciation. To compute the net exports measured in dollar, it is natural to introduce the exchange rate to the fraction of the exports $\xi_t e_t$. Throughout the paper, we simplify our model by setting the export component to $\xi_t = 1$ for $t = 0, 1$. Consequently, ι_t can be interpreted as *net* imports.

4.1.2 Financiers

When global financial markets are imbalanced, there exists an excess supply of dollar versus yen, or vice versa, resulting from trade flows. The financiers are randomly selected from the households of two countries to manage the financial firms. We assume that each fi-

[6] It is equivalent to assume that a small country would not want to launch a trade war with a big country with whom she has trade relationship.

[7] The Japanese counterparts have a symmetric formulation, i.e. $\lambda_t^* = 1$ and $p_{H,t} C_{H,t} = \xi_t^*/(1+\tau_t^g)$.

nancier maximizes the expected value of her firm subject to a credit constraint: $\max V_t = \mathbb{E}_t \left[\beta \left(R - R^* \frac{e_{t+1}}{e_t} \right) \right] q_t$ s.t. $V_t \geq \Gamma_t q_t^2 / e_t$, where q_t is the value of dollar-denominated bonds and the valuation component in the squared bracket corresponds to the households' currency trading. The credit constraint acts as *limited commitment* and we follow closely the specification employed by Maggiori (2017), Gabaix and Maggiori (2015) and Gertler and Kiyotaki (2010).[8] Specifically, we have that

$$\frac{V_t}{e_t} \leq \left| \frac{q_t}{e_t} \right| \left(\Gamma_t \left| \frac{q_t}{e_t} \right| \right) = \Gamma_t \left(\frac{q_t}{e_t} \right)^2,$$

where Γ_t in the first round bracket indicates the portion of dollar-denominated bond value which might be diverted by the financiers. If the financiers divert the funds they intermediate, their firms are unwound and the households can only recover the residual value of financial firms $1 - \Gamma_t \left| \frac{q_t}{e_t} \right|$, where $\Gamma_t = \gamma_t V(e_{t+1})$ with $\gamma_t \geq 0$ captures a limited *risk-bearing capacity* in the financial sector. This formulation highlights the idea that financiers' outside option increase in the size of their balance sheet and also in the volatility of exchange rate, which is affected indirectly by the global trade restrictions.

Similar to Gabaix and Maggiori (2015), the optimal value for the financier to loan dollar-denominated bonds is solved to be $q_t = \frac{1}{\Gamma_t} \mathbb{E} \left[e_t - e_{t+1} \frac{R^*}{R} \right]$. For simplicity, we consider the scenario of equal interest rates across countries $R = R^* = 1$. Integrating this demand function over the unit mass of financiers, we obtain the aggregate demands for dollar-denominated bonds as

$$Q_t = \frac{1}{\Gamma_t} \mathbb{E} \left[e_t - e_{t+1} \right], \tag{5}$$

where Γ_t denotes a *time-varying* risk-bearing capacity for that allows us to explore its endogenous property.

[8] See, among others, Caballero and Krishnamurthy (2001), Kiyotaki and Moore (1997), and Hart and Moore (1994).

4.1.3 Equilibrium Exchange Rates

Aligned with Gabaix and Maggiori (2015), in the long run we assume that the external account will be balanced, i.e. the US export is equal to its import in the terminal period:

$$e_2 = \iota_2 = \bar{\iota}, \tag{6}$$

where $\bar{\iota}$ represents the frictionless US import in the absence of trade restrictions worldwide τ_t^g and of the financial disruptions associated with Γ_t.

Moreover, the demand for dollar versus yen must be cleared in each period. We define the demand function of financiers Q_t and the following market-clearing conditions:

$$\frac{e_0}{1+\tau_0^g} - \iota_0 + Q_0 = 0, \tag{7}$$

$$\frac{e_1}{1+\tau_1^g} - \iota_1 + Q_1 = 0. \tag{8}$$

Different from a two-period model, the financiers only intermediate new flows Q_1 at $t = 1$ while the stock of financial assets Q_0 is held passively (as the long-run investors in this context) by the households until $t = 2$.[9]

Proposition 1. *Assuming the trade restrictions worldwide are persistent with both τ_0^g and $\tau_1^g > 0$, the equilibrium exchange rates follow*

$$e_1 = \frac{\Gamma_1 \iota_1 + \bar{\iota}}{1 + \frac{\Gamma_1}{1+\tau_1^g}}, \quad e_0 = \frac{\left(\frac{\Gamma_1 \iota_1 + \bar{\iota}}{1 + \frac{\Gamma_1}{1+\tau_1^g}}\right)\Gamma_1 + \iota_0 \Gamma_0}{1 + \frac{\Gamma_0}{1+\tau_0^g}}, \tag{9}$$

where e_1 is increasing in τ_1^g and e_0 is increasing in both τ_0^g, τ_1^g.

The impact from the initial action of trade restriction τ_0^g vanishes in the determination of e_1 while the long-lasting trade restriction, expressed by the consecutively rising tariffs τ_1^g, influences both e_0 and e_1. This resilient influence of the trade restrictions worldwide on

[9]See the details in the online appendix of Gabaix and Maggiori (2015).

exchange rates is not obvious in a two-period model, which was useful only in explaining the transmission mechanism of any *transitory* trade restriction on the exchange rate. On the other hand, the τ_1^g in e_0 actually amplifies the degree of currency appreciation caused by τ_0^g and leads to an overshooting effect in e_0. The time-series influences of the trade restrictions worldwide can be demonstrated more clearly in the *returns* of exchange rate.

Proposition 2. *Assuming imperfect financial market $\Gamma_1 > 1 + \tau_0^g$, the equilibrium exchange rate returns $\Delta e_{t+1} = \frac{e_{t+1} - e_t}{e_t}$ for $t = \{0, 1\}$ follow*[10]

$$\Delta e_1 = \frac{\frac{\Gamma_1}{1+\tau_0^g} e_1 - \iota_0 \Gamma_1}{e_1 + \iota_0 \Gamma_1}, \quad \Delta e_2 = \frac{\Gamma_1 \left(\frac{\bar{\iota}}{1+\tau_1^g} - \iota_1 \right)}{\Gamma_1 \iota_1 + \bar{\iota}}, \tag{10}$$

where Δe_1 is increasing in trade restriction τ_1^g but Δe_2 is decreasing in trade restriction τ_1^g.

The assumption of trade restriction worldwide being persistent up to medium term only reflects that the presidency interchanges between political parties are within the investor's expectation. At periods 0 and 1, more trade restrictions are imposed under the Republican presidency. On the other hand, the trade restrictions are relaxed and thus converge to 0 at period 2, as in the long run the party to which the US president belong will eventually change to the other one, i.e. Democratic party.

For the medium-term exchange rate return Δe_1, the *persistent* trade restrictions worldwide, i.e. both τ_0^g and $\tau_1^g > 0$, lead to a currency appreciation in the non-US country. Moving one period forward, knowing that the trade restrictions worldwide will be eased to the frictionless equilibrium eventually, the current trade restrictions in place τ_1^g are considered as *transitory* only. As a result, the long-term exchange rate return Δe_2 is instead depreciating in trade restrictions. To sum up, Proposition 2 shows that the equilibrium currency returns depend on the persistence of trade restrictions worldwide.

As $\partial e_1 / \partial \tau_0^g = 0$, the exchange rate return at period 1 is decreasing in the initial trade restrictions worldwide τ_0^g. On the contrary, the persisting trade restrictions worldwide τ_1^g is positively associated with the exchange rate return when the risk-bearing capacity is

[10] Note that we keep e_1 in Δe_1 for the sake of tractability. By replacing e_1 with (9), the closed-form solution for Δe_1 should consist of only the exogenous variables.

far from limitless. To demonstrate this, we apply the chain rule to the partial derivative $\frac{\partial \Delta e_1}{\partial \tau_1^g} = \frac{\partial \Delta e_1}{\partial e_1} \frac{\partial e_1}{\partial \tau_1^g}$. From (9), we know the sign of the second part as $\partial e_1 / \partial \tau_1^g > 0$. Therefore, the sign of the partial derivative $\frac{\partial \Delta e_1}{\partial \tau_1^g}$ can be determined by the first term:

$$\frac{\partial e_1}{\partial \tau_1^g} = \frac{\iota_0 \Gamma_1 \left(\frac{\Gamma_1}{1+\tau_0^g} - 1 \right)}{\left(e_1 + \iota_0 \Gamma_1\right)^2}. \tag{11}$$

The partial derivative $\frac{\partial \Delta e_1}{\partial \tau_1^g} > 0$ when the risk-bearing capacity deviates from the frictionless level, i.e. the condition of $\Gamma_1 > 1 + \tau_0^g$ holds. Since τ_0^g is bounded by one and a moderate value for relative risk aversion usually ranges between 2 and 3, this inequality does not imply an extreme value of Γ_1 and hence the implied limit on the risk-bearing capacity needs not to match the crisis level.

Proposition 3. *Assuming the persistent trade restrictions worldwide are also sufficiently high,* $\tau_1^g > \frac{\bar{\iota}}{\iota_1} - 1$, *we find*

$$\frac{\partial \Gamma_1}{\partial \tau_1^g} > 0, \tag{12}$$

where the endogenous risk-bearing capacity is worsening in the strengths of trade restrictions worldwide.

To further illustrate the transmission mechanism from the trade restrictions worldwide to the currency market, we explore the sign of the partial derivative of the stochastic risk-bearing capacity in the intermediate term with respect to the trade restrictions worldwide. First, let us express Γ_1 as a function of τ_1^g by rewriting (9):

$$\Gamma_1 = \frac{\bar{\iota} - e_1}{\frac{e_1}{1+\tau_1^g} - \iota_1}, \tag{13}$$

where both e_1 and τ_1^g will affect the Γ_1 given a change in τ_1^g. We apply the chain rule to help us determine the influence of the trade restrictions worldwide τ_1^g: $\partial \Gamma_1 / \partial \tau_1^g = (\partial \Gamma_1 / \partial e_1) \cdot (\partial e_1 / \partial \tau_1^g)$. From Proposition 1, we have already known the second component $\partial e_1 / \partial \tau_1^g > 0$.

What is left to examine is the sign of the first partial derivative:

$$\frac{\partial \Gamma_1}{\partial e_1} = \frac{\iota_1 - \frac{\bar{\iota}}{1+\tau_1^g}}{\left(\frac{e_1}{1+\tau_1^g} - \iota_1\right)^2}. \tag{14}$$

Similar to Proposition 2, we need to impose a condition $\tau_1^g > \frac{\bar{\iota}}{\iota_1} - 1$ to guarantee a positive partial derivative in (14). This condition gives a lower bound for the trade restrictions worldwide τ^g, implying that a significantly stronger trade restrictions at global level are required to cause a negative impact on the financier's risk-bearing capacity. Moreover, the lower bond $\frac{\bar{\iota}}{\iota_1} - 1$ measures essentially the imports *deviation* in percentage from the frictionless counterpart, and hence that highlights again the sensitivity of household's demand with respect to the trade restrictions worldwide plays an important role in the links between trade restrictions and the risk-bearing capacity.

We establish the relationship between the trade policy and the international financial market. This theoretical framework formalize several pieces of anecdote evidence that cast doubts on the influence of the trade policy on the financial market disruptions by incorporating the trade restrictions worldwide (τ^g) in an exchange rate determination model. Note that the main findings in the empirical section confirm the proposed theory. Going forward, we would like to validate our model prediction with an alternative proxy for the risk-bearing capacity in a different empirical setting. Particularly, we want to examine the international market dynamic around the trade policy events by collecting implied volatility data from the currency options and the demand data from the currency futures.

4.1.4 Empirical Extension: Currency Futures Demand

There is a derived model prediction that has not yet been examined in the previous sections. Proposition 3 predicts a negative relationship between risk-bearing capacity and the strengths of trade restrictions worldwide. In order to proxy the risk-bearing capacity in the context of currency market, we extract the demand data for currency futures from the Commodity Futures Trading Commission (CFTC). Garleanu et al. (2009) documented a positive relationship between the risk-bearing capacity of market-makers and the demand pressure of financial derivatives. Combined their finding with Proposition 3 in an application of currency futures, we expect to find a decrease in the futures' demand when the trade restrictions worldwide are anticipated to be elevated. The univariate panel regression in the following is useful for investigating this implication.

$$\text{Currency Futures Demand}_{i,t+1} = \alpha + \beta \; \textit{Trade Policy Index}_t + \gamma_i + \varepsilon_{t+1}, \tag{15}$$

where γ_i is the currency fixed effect. The Trade Policy Index is a low-frequency (monthly) time series constructed from a high-frequency (daily) time series that takes value of 1 or -1 when there was a trade dispute or a trade deal occurring on a specific day, and of 0 in the absence of any type of trade policy event. A list of events used in this paper can be found in Table 7. We then take the sum of all the values within each month. Quantitatively, the Trade Policy Index used in the regression (15) is at monthly frequency and ranges from -1 to 3 in our sample.

This exercise shows a positive and significant estimate of α (≈ 0.03 with a clustered standard error of 0.01) coupled with a negative and significant estimate of β (≈ -0.05 with a clustered standard error of 0.02). This result confirms our expectation that a decreasing demand of currency futures is associated with the trade disputes that can be considered as a proxy for trade restrictions. Furthermore, this result is also aligned with the main finding on the exchange rate returns (See Tables 4 to 6) that the foreign currencies is expected to appreciate under Republican presidencies during which there is on average stronger presence of trade restrictions (see Table 7), leading to the mitigation of the hedging demand on the foreign currencies. For the robustness, we rerun the regression (15) but this time with the demand of US dollar future. Intuitively, we discover a *positive* and significant estimate of β (≈ 0.10 with a clustered standard error of 0.05) while the estimate of α (≈ 0.09 with a clustered standard error of 0.01) is found positive and significant. In addition to the consistent signs on the constant term α, the sign flip on the estimate of βs is completely aligned with our model prediction, implying the following two findings: an anticipated Dollar depreciation associated with a rising hedging demand, and an anticipated appreciation of foreign currencies associated with a falling demand.

4.1.5 Empirical Extension: Customs and Duties

Motivated by the literature of partisanship in political science, we collect the customs and duties data which is a direct measure of tariffs in order to investigate the exact role of tariffs in exchange rate returns between Democratic and Republican presidential cycles. We run panel regressions based on the following specification

$$\Delta s_{i,t+1} = \alpha_i + \beta_1 DP_t + \beta_2 TT_{i,t} + \beta_3 DP_{i,t} \times TT_{i,t} + \gamma' X_t + \varepsilon_{t+1}, \tag{16}$$

where $TT_{i,t}$ denotes trade tariffs (customs and import duties) as percentage of total import for country i as in Lohmann and O'Halloran (1994), X_t is a set control variables, and α_i refers to time-invariant country fixed effects. While β_2 captures the correlation between trade tariffs and exchange rate returns under Republican presidents, the sum of β_2 and β_3 quantifies the impact of trade tariffs under Democratic presidents. We source quarterly data on customs and import duties from the World Bank database and linearly de-trend them as they exhibit a downward trend due to globalization. We then retrieve monthly observations by forward filling. The set of control variables includes the US Federal tax revenue as percentage of GDP (a proxy for US fiscal policy), US import as percentage of GDP (a proxy for US demand of foreign goods and services), and country-level GDPs (a proxy for country size).

Table 9 about here

We report the least-square estimates of our panel regressions in Table 9, while clustering standard errors (reported in parentheses) at the country and time (calendar month) dimension. In specification (1), we find a positive and statistically significant estimate of β_2 (0.015 with a standard error of 0.004) as well as a negative and statistically significant estimate of β_3 (0.024 with a standard error of 0.003). Taken together, these estimates imply a weaker US dollar under Republican presidents since $\beta_2 > 0$ (recall that a positive exchange rate return means US dollar depreciation) and a stronger US dollar under Democratic presidents since $\beta_2 + \beta_3 < 0$ with respect to trade tariffs. Also, the estimate of β_1 associated with the presidential dummy becomes weakly significant, thus suggesting that part of its explanatory power is indeed captured by the role of trade policies. In specification (3), we control for the US fiscal policy as well as its interaction with the presidential dummy and uncover no qualitatively difference in our estimates of β_1 and β_2. Our findings seem to differ from the relationship between stock returns under different presidential regimes and fiscal policy highlighted by (Pástor and Veronesi, 2020). In specification (5), we further control for the US demand of foreign goods, whose coefficient estimate is positive and statistically significant. Moreover, while β_2 and β_3 remain highly statistically significant, the estimate of β_1 becomes statistically insignificant. Overall, trade policies measured via the impact of trade tariffs appear to play an important role to explain the exchange rate return difference between Democratic and Republican presidential terms.

4.2 Model's Discussion

Our model builds on Gabaix and Maggiori (2015) with the exchange rate dynamic depending on the tariff differential. The credit constraint of financier results in an imperfect currency market which exhibits a pattern of foreign exchange returns aligned with the occurrence of tariff differential. The novelty here is the role of the trade policy differential beyond the financial disruptions. The model predictions are supported by the empirical evidence and therefore rationalize the stylized fact of foreign exchange return difference associated with the presidential cycle.

However, this model setup does have certain limitations. For instance, our work abstracts away from the assumption of recursive preferences with long-run risks (Colacito et al., 2018) and therefore the implications of country-specific exposures to global growth news shock are beyond our discussion. In addition, our model differs from Liao and Zhang (2021) in the roles of financier. Their financier produces a financial derivative (i.e. the forward) in order to satisfy investor's hedging demand while our financier only absorbs the excessive demand from the households. Our model does not generate implications on the forward contract.

5 Conclusions

We study the relationship between exchange rate returns and US presidential cycles. Empirically, we document an average US dollar appreciation of 4.31% per annum during Democratic presidencies but a depreciation of 1.25% per annum during Republican presidential terms. The difference between these average exchange rate returns amounts to 5.56% per annum, which is both economically and statistically significant. Several possible explanations, including interest rate differentials, inflation rate differentials, real business cycles, and foreign political cycles, have been ruled out.

As a further investigation, we study the role of trade policy implemented by the US presidents. We first show high tariffs and US dollar depreciation are correlated. Then, we find that trade policy events can be translated into risks for foreign exchange markets. These findings are not driven by country-specific, trade-related characteristics such as country size and distance. Additionally, we extend the exchange rate determination model of Gabaix and Maggiori (2015) to rationalize the empirical findings. In this model, trade tariffs leads to financial disruptions due to the limited risk-bearing capacity of financiers who intermediate the global demand for currencies. The model prediction of higher volatilities in exchange rate matches our empirical findings. Future works can extend our results to better understand the US presidential cycle in foreign exchange markets. For instance, investigating the characteristics of political institution such Congress or Parliament would provide an additional dimension of cross-country variations.

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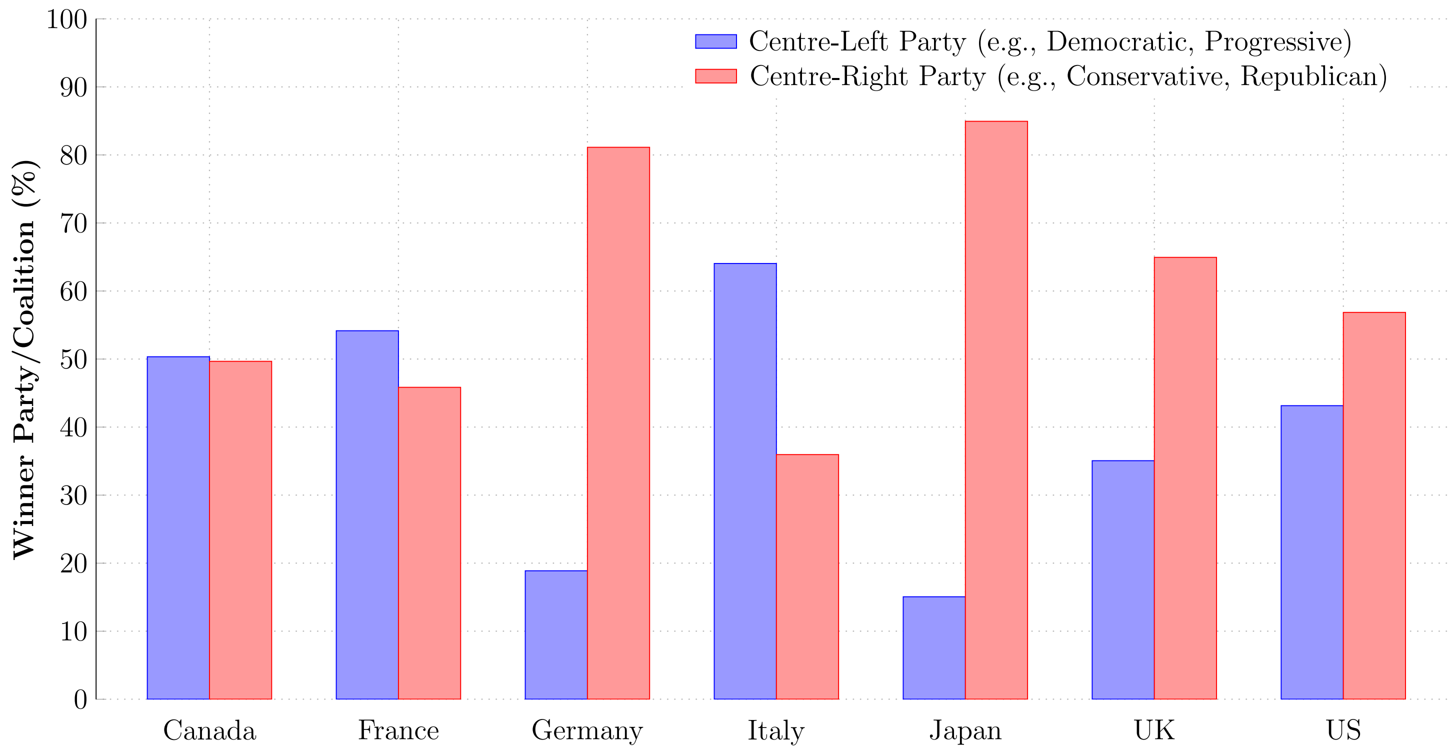

Figure 1. Political Cycles of G7 Countries

This figure displays the political cycle in each of the G7 country. The sample runs from October 1983 to October 2020.

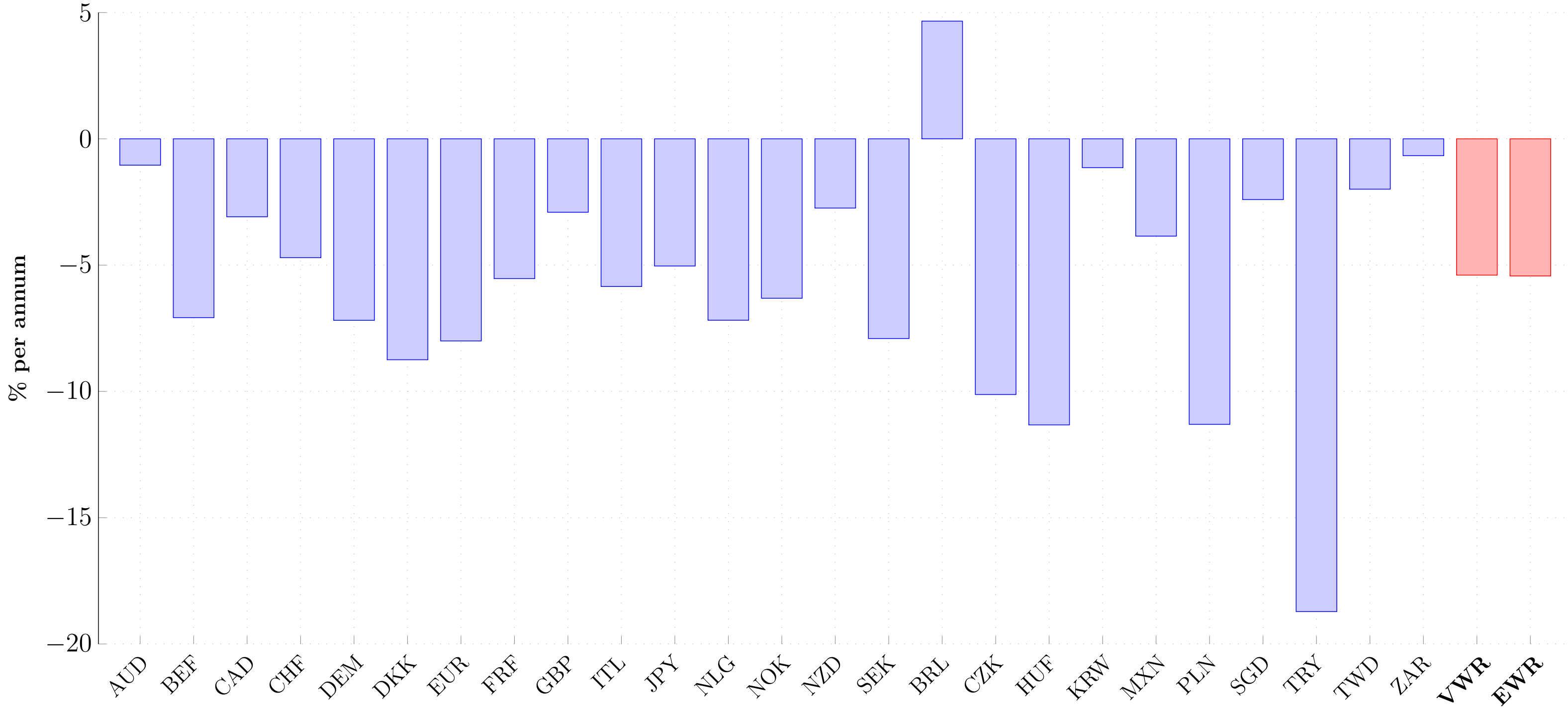

Figure 2. US Political Cycles and Exchange Rate Returns

This figure displays the difference in average exchange rate returns in percentage per annum between Democratic and Republican presidential terms. VWR is a basked of GDP-weighted returns, whereas EWR is a basked of equally-weighted returns. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream whereas GDP data are from the World Economic Outlook Database.

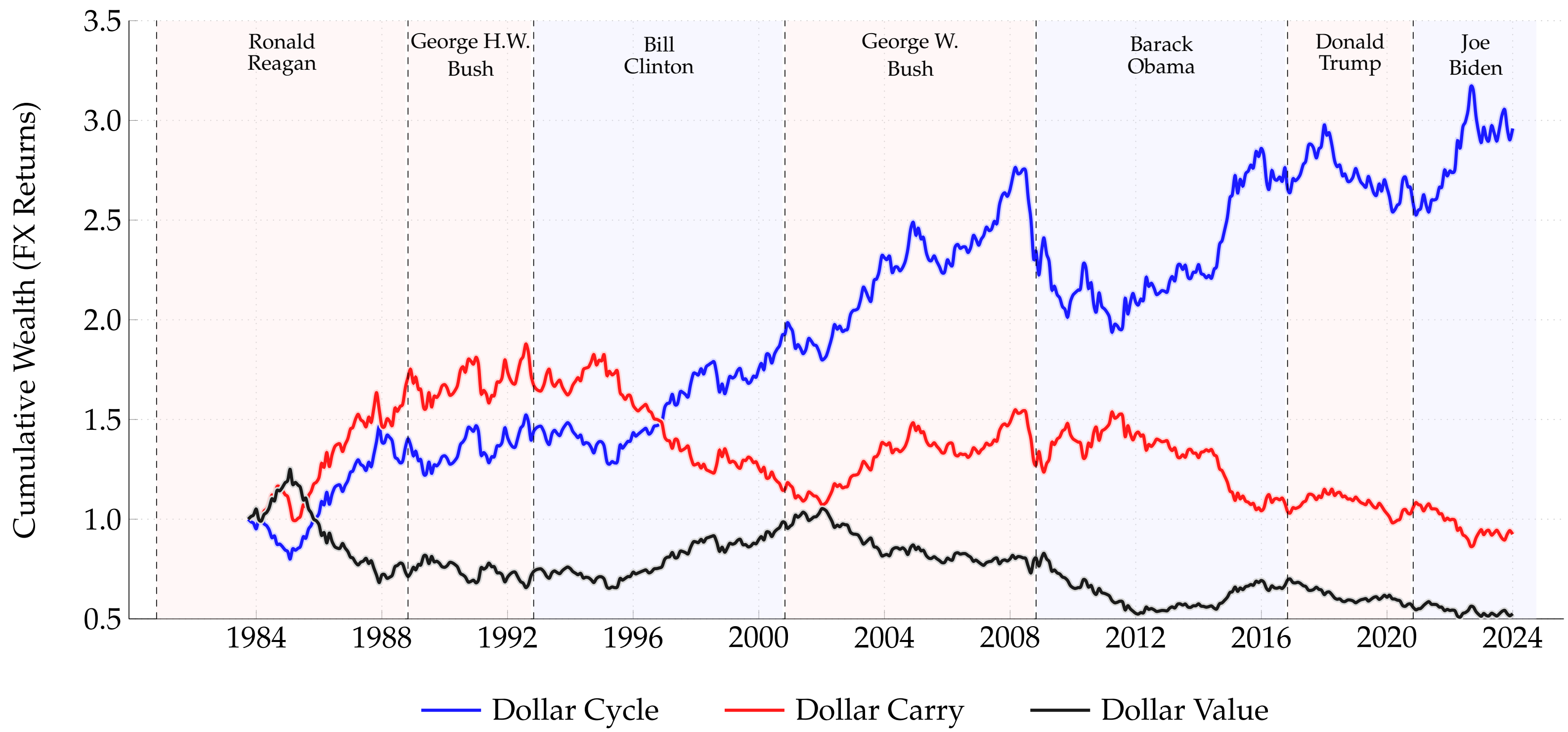

Figure 3. Dollar-based Strategies: GDP-weighted Exchange Rate Returns

This figure displays, for equally-weighted dollar-based trading strategies, the cumulative wealth based on exchange rate returns. The dollar cycle buys (sells) the US dollar and sells (buys) a GDP-weighted basket of foreign currencies during Democratic (Republican) presidential terms. The dollar carry buys (sells) the US dollar and sells (buys) a GDP-weighted basket of foreign currencies whenever the short-term US interest rate is above the average foreign short-term interest rate. Interest rate differentials are implied from spot and forward exchange rates. The dollar value buys (sells) the US dollar and sells (buys) a GDP-weighted basket of foreign currencies whenever the US inflation rate is below the average foreign inflation rate. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Spot and one-month forward exchange rates (monthly frequency), and year-on-year inflation rates (monthly) are from Datastream, whereas GDP data (yearly frequency) are from the World Economic Outlook Database.

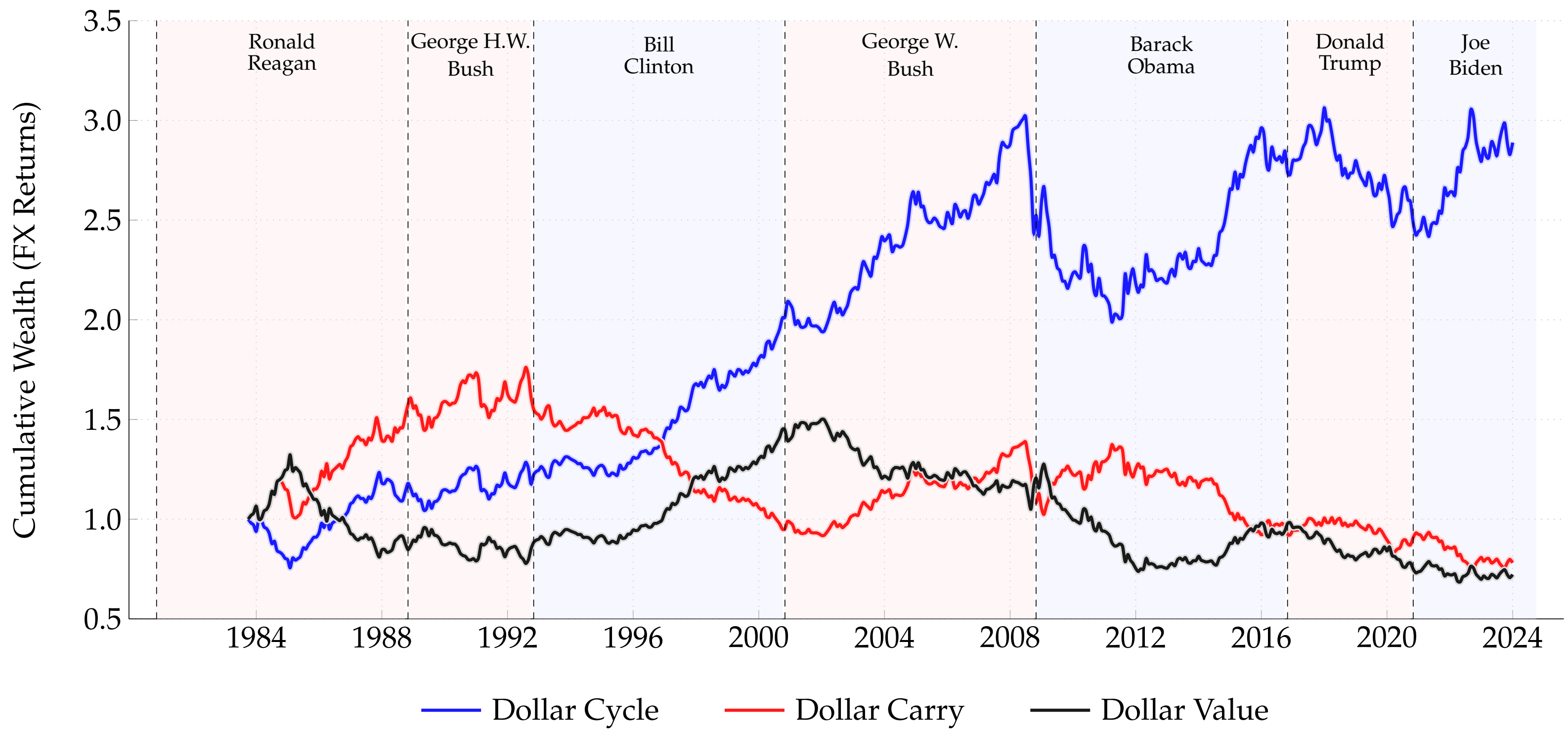

Figure 4. Dollar-based Strategies: Equally-weighted Exchange Rate Returns

This figure displays, for equally-weighted dollar-based trading strategies, the cumulative wealth based on exchange rate returns. The dollar cycle buys (sells) the US dollar and sells (buys) an equally-weighted basket of foreign currencies during Democratic (Republican) presidential terms. The dollar carry buys (sells) the US dollar and sells (buys) an equally-weighted basket of foreign currencies whenever the short-term US interest rate is above the average foreign short-term interest rate. Interest rate differentials are implied from spot and forward exchange rates. The dollar value buys (sells) the US dollar and sells (buys) an equally-weighted basket of foreign currencies whenever the US inflation rate is below the average foreign inflation rate. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Spot and one-month forward exchange rates (monthly frequency), and year-on-year inflation rates (monthly) are from Datastream.

Table 1. Summary Statistics: Exchange Rate Returns

This table presents means and standard deviations in percentage per annum of country-level nominal exchange rate returns. DP denotes the Democratic Presidential terms, RP indicates the Republican Presidential terms, and DP-RP is the corresponding difference between means and standard deviations. VWR is a basked of GDP-weighted returns, whereas EWR is a basked of equally-weighted returns. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Spot exchange rates (monthly frequency) are from Datastream, whereas GDP data (yearly frequency) are from the World Economic Outlook Database.

	Full Sample		Democratic (DP)		Republican (RP)		DP – RP	
	mean	std	mean	std	mean	std	mean_{dif}	std_{dif}
AUD	–0.57	11.67	–1.10	11.36	–0.05	11.98	–1.04	–0.62
BEF	2.90	11.62	–1.30	9.24	5.78	12.97	–7.08	–3.73
CAD	–0.03	7.34	–1.59	7.46	1.49	7.21	–3.09	0.25
CHF	2.27	11.00	–0.18	10.82	4.52	11.14	–4.71	–0.32
DEM	3.01	11.75	–1.25	9.30	5.93	13.14	–7.19	–3.84
DKK	1.27	10.15	–3.16	9.69	5.58	10.46	–8.75	–0.76
EUR	–0.31	9.41	–4.14	9.97	3.86	8.62	–8.00	1.35
FRF	2.37	11.20	–0.91	9.06	4.62	12.46	–5.53	–3.40
GBP	–0.40	9.89	–1.92	8.48	0.99	11.02	–2.91	–2.54
ITL	–0.21	11.33	–3.68	9.12	2.17	12.61	–5.85	–3.49
JPY	1.17	10.77	–1.46	11.47	3.58	10.06	–5.03	1.41
NLG	2.97	11.68	–1.29	9.32	5.89	13.02	–7.18	–3.70
NOK	–0.36	11.34	–3.56	11.15	2.75	11.48	–6.31	–0.33
NZD	0.66	12.17	–0.73	12.32	2.01	12.03	–2.74	0.28
SEK	–0.36	10.96	–4.37	11.39	3.54	10.43	–7.91	0.96
BRL	–3.48	17.38	–1.24	15.67	–5.90	19.08	4.66	–3.40
CZK	0.69	11.33	–3.26	11.45	6.86	10.93	–10.12	0.52
HUF	–4.17	12.53	–8.69	12.73	2.64	12.01	–11.33	0.72
KRW	–1.67	13.33	–2.13	15.78	–0.98	8.42	–1.14	7.36
MXN	–5.68	13.73	–7.22	15.24	–3.37	11.08	–3.85	4.16
PLN	–2.64	12.38	–7.07	12.48	4.23	11.98	–11.30	0.50
SGD	1.25	5.39	0.03	6.08	2.43	4.61	–2.40	1.47
TRY	–20.99	16.47	–28.31	14.26	–9.59	19.02	–18.72	–4.77
TWD	–0.54	5.27	–1.33	5.81	0.66	4.34	–1.99	1.47
ZAR	–6.86	15.29	–7.21	12.97	–6.55	17.17	–0.67	–4.20
VWR	–0.32	8.17	–3.15	7.73	2.28	8.50	–5.43	–0.77
EWR	–1.33	8.17	–4.15	7.85	1.25	8.40	–5.40	–0.55

Table 2. Summary Statistics: Currency Excess Returns

This table presents means and standard deviations in percentage per annum of country-level nominal currency excess returns. DP denotes the Democratic Presidential terms, RP indicates the Republican Presidential terms, and DP-RP is the corresponding difference between means and standard deviations. VWR is a basked of GDP-weighted returns, whereas EWR is a basked of equally-weighted returns. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Spot and one-month forward exchange rates (monthly frequency) are from Datastream, whereas GDP data (yearly frequency) are from the World Economic Outlook Database.

	Full Sample		Democratic (DP)		Republican (RP)		DP – RP	
	mean	std	mean	std	mean	std	mean_{dif}	std_{dif}
AUD	1.88	11.74	0.34	11.41	3.37	12.06	−3.03	−0.66
BEF	3.59	11.67	−1.41	9.25	7.01	13.02	−8.42	−3.77
CAD	0.45	7.37	−1.55	7.46	2.39	7.25	−3.93	0.21
CHF	0.32	11.02	−1.91	10.80	2.36	11.20	−4.27	−0.40
DEM	2.03	11.73	−1.49	9.23	4.45	13.17	−5.94	−3.94
DKK	1.51	10.22	−3.14	9.72	6.02	10.53	−9.16	−0.81
EUR	−1.12	9.45	−4.95	9.99	3.06	8.69	−8.01	1.29
FRF	3.85	11.24	−0.32	9.02	6.71	12.51	−7.03	−3.49
GBP	0.81	9.95	−1.52	8.47	2.96	11.12	−4.48	−2.65
ITL	3.74	11.30	−0.17	8.83	6.54	12.76	−6.71	−3.92
JPY	−1.25	10.81	−3.81	11.47	1.09	10.15	−4.90	1.32
NLG	2.22	11.72	−1.73	9.23	4.93	13.14	−6.66	−3.92
NOK	1.28	11.37	−2.71	11.12	5.15	11.52	−7.86	−0.40
NZD	4.09	12.34	1.20	12.32	6.89	12.33	−5.70	0.00
SEK	0.46	11.01	−4.03	11.38	4.82	10.51	−8.85	0.87
BRL	5.15	15.39	6.08	15.20	3.92	15.70	2.16	−0.50
CZK	1.36	11.81	−2.48	12.37	6.19	10.95	−8.68	1.42
HUF	2.14	13.25	−1.03	14.01	5.90	12.24	−6.93	1.77
KRW	0.51	10.75	0.37	12.66	0.65	8.33	−0.28	4.33
MXN	3.92	10.86	5.24	10.64	2.26	11.16	2.97	−0.52
PLN	2.91	12.83	−0.33	13.32	7.07	12.12	−7.41	1.19
SGD	0.27	5.41	−0.56	6.10	1.08	4.64	−1.64	1.45
TRY	5.03	16.39	1.33	13.38	10.15	19.78	−8.82	−6.41
TWD	−1.58	5.39	−2.11	6.12	−0.90	4.32	−1.21	1.80
ZAR	0.33	15.26	−0.35	13.00	0.94	17.09	−1.29	−4.08
VWR	0.25	8.27	−2.44	7.85	2.72	8.60	−5.16	−0.75
EWR	1.38	8.25	−0.93	7.89	3.50	8.53	−4.43	−0.64

Table 3. Summary Statistics: Real Exchange Rate Returns

This table presents means and standard deviations in percentage per annum of country-level real exchange rate returns. DP denotes the Democratic Presidential terms, RP indicates the Republican Presidential terms, and DP-RP is the corresponding difference between means and standard deviations. VWR is a basked of GDP-weighted returns, whereas EWR is a basked of equally-weighted returns. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Spot exchange rates (monthly frequency) and year-on-year inflation rates (monthly frequency) are from Datastream, whereas GDP data (yearly frequency) are from the World Economic Outlook Database.

	Full Sample		Democratic (DP)		Republican (RP)		DP – RP	
	mean	std	mean	std	mean	std	mean_{dif}	std_{dif}
AUD	0.03	11.73	–1.01	11.42	1.03	12.04	–2.04	–0.62
BEF	2.19	11.63	–1.95	9.26	5.02	12.99	–6.97	–3.73
CAD	–0.35	7.38	–2.10	7.48	1.34	7.25	–3.44	0.23
CHF	0.72	11.04	–2.00	10.86	3.21	11.17	–5.22	–0.30
DEM	1.91	11.76	–1.57	9.31	4.30	13.18	–5.87	–3.87
DKK	0.75	10.21	–3.57	9.76	4.95	10.51	–8.52	–0.75
EUR	–1.22	9.49	–5.39	10.07	3.32	8.66	–8.71	1.41
FRF	2.01	11.23	–1.88	9.08	4.68	12.47	–6.57	–3.39
GBP	–0.16	9.94	–1.77	8.54	1.31	11.06	–3.08	–2.52
ITL	1.97	11.36	–2.54	9.14	5.06	12.62	–7.59	–3.48
JPY	–0.98	10.80	–3.44	11.51	1.28	10.08	–4.72	1.43
NLG	1.57	11.68	–1.60	9.37	3.74	13.04	–5.35	–3.67
NOK	–0.21	11.39	–3.54	11.20	3.02	11.52	–6.56	–0.32
NZD	1.41	12.33	–0.97	12.37	3.71	12.27	–4.68	0.10
SEK	–0.43	11.04	–5.01	11.45	4.01	10.49	–9.01	0.96
BRL	0.27	17.42	2.88	15.63	–2.53	19.18	5.41	–3.55
CZK	2.73	11.40	–0.06	11.63	7.09	10.96	–7.15	0.67
HUF	1.24	12.59	–1.29	12.85	5.06	12.16	–6.35	0.69
KRW	–1.25	13.36	–1.40	15.82	–1.03	8.44	–0.37	7.39
MXN	–0.13	13.82	0.65	15.37	–1.30	11.10	1.96	4.27
PLN	1.92	12.40	0.39	12.64	4.29	12.03	–3.90	0.61
SGD	0.17	5.43	–0.34	6.15	0.67	4.64	–1.01	1.51
TRY	5.64	16.48	7.12	14.21	3.33	19.54	3.78	–5.33
TWD	–1.71	5.34	–2.25	5.89	–0.91	4.38	–1.34	1.52
ZAR	–1.83	15.39	–3.24	13.07	–0.55	17.27	–2.69	–4.20
VWR	–0.38	8.20	–2.91	7.78	1.94	8.52	–4.85	–0.74
EWR	0.31	8.19	–1.56	7.88	2.03	8.44	–3.59	–0.55

Table 4. Exchange Rate Returns and Presidential Cycles

This table presents estimates of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). A presidential cycle starts in November when the elections take place and ends four years after in October. Panel A presents estimates from pooled regressions whereas Panel B from panel regressions with currency fixed effects. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream.

	Panel A: Pooled Regressions						Panel B: Fixed Effects Regressions					
	α		DP		R^2 (%)	N	α		DP		R^2 (%)	N
All Countries	1.645	(1.874)	−5.775**	(2.628)	0.501	9,240	1.398	(1.749)	−5.294**	(2.540)	1.423	9,240
Remove AUD	1.740	(1.899)	−6.025**	(2.637)	0.545	8,771	1.482	(1.758)	−5.525**	(2.541)	1.523	8,771
Remove BEF	1.543	(1.858)	−5.718**	(2.623)	0.490	9,058	1.306	(1.734)	−5.259**	(2.539)	1.426	9,058
Remove CAD	1.653	(1.938)	−5.913**	(2.698)	0.508	8,771	1.396	(1.799)	−5.414**	(2.605)	1.446	8,771
Remove CHF	1.474	(1.875)	−5.806**	(2.637)	0.502	8,757	1.227	(1.747)	−5.327**	(2.545)	1.439	8,757
Remove DEM	1.539	(1.858)	−5.715**	(2.623)	0.490	9,058	1.303	(1.734)	−5.257**	(2.539)	1.426	9,058
Remove DKK	1.425	(1.859)	−5.604**	(2.621)	0.464	8,771	1.168	(1.732)	−5.106*	(2.526)	1.413	8,771
Remove EUR	1.571	(1.888)	−5.701**	(2.634)	0.481	8,939	1.315	(1.754)	−5.201*	(2.539)	1.431	8,939
Remove FRF	1.571	(1.862)	−5.753**	(2.626)	0.496	9,058	1.333	(1.736)	−5.289**	(2.542)	1.434	9,058
Remove GBP	1.683	(1.886)	−5.927**	(2.660)	0.519	8,757	1.426	(1.748)	−5.428**	(2.570)	1.481	8,757
Remove ITL	1.632	(1.867)	−5.769**	(2.633)	0.499	9,058	1.381	(1.739)	−5.283**	(2.548)	1.449	9,058
Remove JPY	1.530	(1.930)	−5.797**	(2.705)	0.500	8,757	1.278	(1.798)	−5.308*	(2.615)	1.456	8,757
Remove NLG	1.540	(1.858)	−5.715**	(2.623)	0.490	9,058	1.304	(1.734)	−5.257**	(2.539)	1.426	9,058
Remove NOK	1.583	(1.865)	−5.742**	(2.617)	0.493	8,771	1.324	(1.726)	−5.239**	(2.520)	1.471	8,771
Remove NZD	1.624	(1.903)	−5.929**	(2.637)	0.530	8,771	1.369	(1.762)	−5.433**	(2.541)	1.505	8,771
Remove SEK	1.539	(1.869)	−5.657**	(2.619)	0.476	8,771	1.279	(1.731)	−5.152*	(2.521)	1.451	8,771
Remove BRL	1.894	(1.868)	−6.122**	(2.605)	0.588	8,941	1.643	(1.742)	−5.632**	(2.507)	1.590	8,941
Remove CZK	1.472	(1.872)	−5.646**	(2.624)	0.477	8,871	1.194	(1.740)	−5.100*	(2.520)	1.432	8,871
Remove HUF	1.612	(1.871)	−5.524**	(2.604)	0.460	8,879	1.373	(1.727)	−5.055*	(2.505)	1.428	8,879
Remove KRW	1.732	(1.905)	−5.958**	(2.654)	0.540	8,879	1.477	(1.774)	−5.458**	(2.560)	1.524	8,879
Remove MXN	1.811	(1.894)	−5.793**	(2.682)	0.512	8,879	1.585	(1.769)	−5.351**	(2.581)	1.464	8,879
Remove PLN	1.559	(1.871)	−5.544**	(2.605)	0.463	8,873	1.309	(1.729)	−5.053*	(2.504)	1.443	8,873
Remove SGD	1.601	(1.941)	−5.944**	(2.698)	0.508	8,871	1.347	(1.804)	−5.451**	(2.606)	1.420	8,871
Remove TRY	1.982	(1.857)	−5.025*	(2.538)	0.395	8,905	1.870	(1.736)	−4.806*	(2.504)	0.504	8,905
Remove TWD	1.677	(1.921)	−5.942**	(2.695)	0.513	8,879	1.414	(1.793)	−5.425**	(2.601)	1.440	8,879
Remove ZAR	2.131	(1.812)	−6.103**	(2.617)	0.584	8,757	1.847	(1.737)	−5.553**	(2.526)	1.494	8,757

Table 5. Controlling for Local Political Cycles

This table presents estimates, for pooled regressions, of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US and/or a dummy variable FC that takes on the value of one (zero) during Centre-Left (Centre-Right) Political terms in major foreign countries like Canada, France, Germany, Italy, Japan, and the UK. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). The US presidential cycle starts in November when the elections take place and ends four years after in October. The election dates and the duration of local political cycle vary across G7 countries. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample consists of monthly observations between October 1983 and January 2024 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
DP	−5.775**							−5.600**	−5.793**	−5.762**	−5.843**	−6.768**	−5.649**	−6.040**
	(2.586)							(2.582)	(2.569)	(2.600)	(2.555)	(2.624)	(2.634)	(2.866)
Canada		−2.367						−1.820						−0.969
		(2.716)						(2.693)						(3.735)
France			0.031						−0.312					3.517
			(2.506)						(2.446)					(4.390)
Germany				0.657						0.386				1.061
				(2.633)						(2.598)				(3.753)
Italy					−0.584						−0.987			−0.134
					(2.488)						(2.447)			(3.439)
Japan						2.279						5.885		7.019
						(5.715)						(5.869)		(7.407)
UK							1.789						0.794	3.727
							(2.538)						(2.552)	(5.061)
α	1.645	0.006	−1.336	−1.479	−1.010	−1.517	−1.933	2.574	1.825	1.545	2.203	1.645	1.308	−1.666
	(1.845)	(2.244)	(2.054)	(1.690)	(2.096)	(1.524)	(1.809)	(2.454)	(2.231)	(2.083)	(2.269)	(1.845)	(2.219)	(5.991)
R^2 (%)	0.501	0.074	−0.011	−0.006	−0.006	0.014	0.033	0.539	0.491	0.491	0.505	0.643	0.498	0.675
N	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240	9,240

Table 6. Controlling for US Business Cycle Fluctuations

This table presents pooled regression estimates of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US while controlling for the term spread (TSP), default spread (DSP), relative interest rate (RR), and the log dividend-price ratio (LDP). Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). The control variables are demeaned and lagged relative to the exchange rate returns. A presidential cycle starts in November when the elections take place and ends four years after in October. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. Exchange rate returns are expressed in percentage per annum. The sample consists of monthly observations between October 1983 and October 2020 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream, the dividend-price ratio is from Robert Shiller's website, and the other data are from the FRED database.

	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
	Panel A: Controls lagged by one month					Panel B: Controls lagged by three months				
DP	−6.222**	−5.571**	−5.677**	−5.399**	−6.244**	−6.436**	−5.433**	−5.657**	−5.546**	−6.513**
	(2.675)	(2.743)	(2.684)	(2.650)	(2.802)	(2.659)	(2.693)	(2.684)	(2.633)	(2.711)
TSP	1.348				1.087	1.690				1.321
	(1.172)				(1.255)	(1.180)				(1.254)
DSP		−5.595			−5.424		−7.766			−8.810
		(5.922)			(6.250)		(5.255)			(5.515)
RR			−0.878		1.948			−1.180		2.946
			(2.049)		(2.072)			(2.009)		(1.970)
LDP				4.328	2.750				3.448	−0.460
				(4.127)	(4.425)				(4.056)	(4.448)
α	1.873	1.571	1.565	1.567	2.129	1.982	1.461	1.554	1.606	2.040
	(1.858)	(1.842)	(1.795)	(1.804)	(1.894)	(1.860)	(1.843)	(1.812)	(1.809)	(1.919)
R^2 (%)	0.675	0.854	0.553	0.652	1.039	0.753	1.118	0.570	0.610	1.345
N	9,348	8,960	9,348	9,348	8,960	9,348	8,924	9,348	9,348	8,924

(continued)

Table 6. Controlling for US Business Cycle Fluctuations *(continued)*

	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
	Panel C: Controls lagged by six months					Panel D: Controls lagged by one year				
DP	−6.495**	−5.436**	−5.612**	−5.576**	−6.053**	−6.995**	−5.391**	−6.028**	−5.730**	−6.605**
	(2.674)	(2.631)	(2.649)	(2.614)	(2.674)	(2.679)	(2.669)	(2.642)	(2.625)	(2.749)
TSP	1.456				0.804	2.245				1.790
	(1.196)				(1.340)	(1.152)				(1.321)
DSP		−9.285			−9.962		−5.181			−4.011
		(3.618)			(3.861)		(3.341)			(3.697)
RR			−2.364		1.898			−1.747		−0.365
			(2.484)		(2.686)			(2.221)		(2.683)
LDP				3.882	−0.202				3.190	−0.943
				(3.894)	(4.349)				(3.710)	(4.368)
α	1.992	1.437	1.549	1.632	1.770	2.227	1.249	1.766	1.682	1.956
	(1.881)	(1.822)	(1.806)	(1.814)	(1.927)	(1.832)	(1.837)	(1.843)	(1.823)	(1.999)
R^2 (%)	0.693	1.377	0.682	0.634	1.449	0.902	0.741	0.618	0.606	0.923
N	9,348	8,870	9,348	9,348	8,870	9,348	8,762	9,348	9,348	8,762

Table 7. List of Trade Policy Events

This table presents a list of 41 global trade events from the year of 1983 to 2020. The superscripts † indicates a trade policy deal.

No.	Date	Trade Policy Events	Democrats	Republicans
1	08/09/1985	Protectionist legislation talks		✓
2	15/04/1987	Trade sanctions on Japan (anti-dumping)		✓
3	01/02/1988	Presidential campaign - oil import fee discussions		✓
4	04/11/1992	Trade war with Europe over farm subsidies	✓	
5	01/02/1993	Oil import fee discussed; Clinton takes office	✓	
6	20/11/1993	Congress passes NAFTA†	✓	
7	30/11/1994	Uncertainty in Senate over passing GATT	✓	
8	23/06/1995	Tariff threat on Japanese autos	✓	
9	30/11/1999	Seattle Protests WTO	✓	
10	11/12/2001	China becomes WTO member†		✓
11	20/03/2002	Bush's Steel tariff in effect		✓
12	03/11/2003	WTO Penalizes steel tariffs†		✓
13	20/03/2006	2006/333/EC EU vs. USA over common wheat†		✓
14	12/10/2006	Softwood Lumber Agreement taking effect†		✓
15	03/01/2007	EU rules for quotas taking effect		✓
16	22/02/2007	2007/444/EC with Canada over common wheat†		✓
17	01/04/2008	India, China, Vietnam and Egypt impose export ban on rice		✓
18	26/06/2008	EU extends the suspension on cereal tariff†		✓
19	27/10/2008	EU reintroduces tariff on cereal taking effect		✓
20	24/09/2014	Canada-European Trade Agreement (CETA) negotiation concludes†	✓	
21	07/05/2015	Brexit - announcement	✓	
22	12/10/2015	Softwood Lumber Agreement ended, extended one year†	✓	
23	17/12/2015	EU Referendum Act, Brexit†	✓	
24	04/02/2016	Trans-Pacific Partnership (TPP) Agreement singed†	✓	
25	22/02/2016	Official EU referendum announced	✓	
26	24/06/2016	Brexit referendum result	✓	
27	12/10/2016	Softwood agreement expires	✓	
28	31/10/2016	CETA signed†	✓	
29	03/01/2017	Tariff threat on Mexico; Trump takes office		✓
30	23/01/2017	US withdrawal from TPP		✓
31	26/01/2017	Threat to Mexico 20% import tariff		✓
32	20/04/2017	Trump signs for the steel investigation		✓
33	03/07/2017	Republican tax plan - import tax debate		✓
34	21/09/2017	CETA enters into forces†		✓
35	22/01/2018	Tariff on washing machine and solar panel B		✓
36	29/05/2018	Trump announces Chinese tariff		✓
37	02/07/2018	Tariffs on Chinese goods take effect		✓
38	31/05/2019	Trump threat Mexico with 5% import tax		✓
39	03/18/2020	US impose tariff on wine food aircraft parts taking effect		✓
40	03/24/2020	Vietnam and Serbia restriction on rice export		✓
41	01/04/2020	Russia restriction on rice export		✓
		Sum of trade policy disputes	10	19
		Sum of trade policy deals	5	7

Table 8. Presidential Cycles and Trade Restrictions

This table presents panel regressions estimates of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US, $MATR$ based on Measure of Aggregate Trade Restrictions, Changes in trade frictions estimated by cross-country trade variables, and the control variables, including US Fiscal Policy (Federal tax revenue) standardized by GDP and Monetary Policy difference across countries. Annual observations on the MATR is applied forward-filled log transformation and first difference. Standard errors (in parentheses) are clustered by currency and time (calendar year) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample consists of monthly observations between October 1983 and October 2020 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream, trade restrictions, quarterly total imports and exports are from the IMF, and other data from the FRED database.

	(1)	(2)	(3)	(4)	(5)	(6)
DP	−5.724*	1.573	2.013	1.899	2.438	−2.485
	(2.910)	(3.967)	(4.084)	(4.067)	(4.385)	(3.349)
$MATR$	0.526**	0.498**	0.614***	0.615***	0.552***	0.563***
	(0.201)	(0.192)	(0.174)	(0.174)	(0.191)	(0.194)
Change in trade frictions (USA Import)		24.247	26.028	26.028	29.887*	
		(14.951)	(15.255)	(15.297)	(16.574)	
Change in trade frictions (USA Export)		−54.478**	−54.311**	−54.032**	−62.041**	
		(19.526)	(20.394)	(20.364)	(22.679)	
Change in trade frictions (Import)			−26.665***	−26.610***	−26.516**	
			(8.746)	(8.666)	(9.892)	
Change in trade frictions (Export)			24.429**	24.275**	26.244**	
			(10.115)	(10.025)	(11.197)	
US Fiscal Policy				0.791	0.885	1.514
				(2.333)	(2.614)	(2.729)
Monetary Policy					−0.722	−0.784
					(0.755)	(0.758)
Change in trade frictions (USA Net Import)						24.952
						(15.704)
Change in trade frictions (Net Import)						−28.496**
						(11.487)
α	1.551	29.005**	28.986**	28.855**	30.714**	−0.231
	(2.031)	(13.004)	(12.748)	(12.631)	(13.267)	(2.091)
R^2 (%)	1.513	2.458	2.769	2.761	2.912	2.158
N	8,138	8,138	7,764	7,764	6,962	6,962

Table 9. Presidential Cycles and Trade Tariffs

This table presents panel regressions estimates of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US, tariffs (custom and import duties) standardized by total import, US tax (Federal tax revenue) standardized by GDP, and the control variables, including the total imports standardized by GDP, the country-level GDP, and VIX. Quarterly observations on trade tariff are detrended. Monthly for trade tariff and GDP are retrieved from quarterly observations via forward filling. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample consists of monthly observations between October 1983 and October 2020 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream, trade tariffs from the World Bank, monthly total imports are from the IMF, VIX is from the CBOE, and other data from the FRED database.

	(1)	(2)	(3)	(4)	(5)	(6)
DP	−4.694	−5.780*	−4.904	−4.028	−4.454	−3.473
	(3.028)	(2.873)	(3.141)	(3.524)	(3.446)	(3.674)
Tariffs	0.016***		0.016***	0.017***		0.017***
	(0.004)		(0.005)	(0.004)		(0.005)
DP × *Tariffs*	−0.024***		−0.024***	−0.024***		−0.023***
	(0.003)		(0.006)	(0.007)		(0.008)
Imports	4.067**		3.993**	3.567*		3.924*
	(1.907)		(1.797)	(1.970)		(2.109)
US Fed Tax		1.806	1.878		−1.075	−0.593
		(2.890)	(2.934)		(2.974)	(3.197)
DP × *US Fed Tax*		−0.547	−1.135		−5.125	−5.277
		(4.915)	(4.035)		(6.845)	(7.117)
US GDP		0.000	0.000		0.000	0.000
		(0.000)	(0.000)		(0.000)	(0.000)
VIX				−0.536	−0.693**	−0.602
				(0.343)	(0.291)	(0.369)
α	−9.456*	2.269	−9.013	1.138	15.468**	2.387
	(4.913)	(3.029)	(5.256)	(8.394)	(6.935)	(9.807)
R^2 (%)	2.146	1.308	2.123	2.763	2.687	2.764
N	4,506	8,422	4,492	3,679	7,354	3,679
VIX Control				✓	✓	✓

Internet Appendix to

"Presidential Cycles and Exchange Rates"

(not for publication)

Abstract

We present supplementary results not included in the main body of the paper.

A The Determination of Real Exchange Rate without Financiers

To illustrate the main finding of our empirical paper, a simple production-based trade model without financial intermediary can serve the purpose.

Consider the tradable goods produced by the US as $C_T = a_T L_T$ where a_T is the technology and L_T is the labor hours for the sector that produces the tradable goods. Denote the price of US tradable goods as P_T and the wage as w. The sector of non-tradable goods is indicated by the subscription of N. The US price index is $P = \phi(P_T, P_N)$ where ϕ is homogeneous of degree 1.

The Japanese variables have similar expressions and is marked with an asterisk. We assume that law of one price holds for tradable goods $P_T = EP_T^*$ but not for the non-tradable goods $P_N \neq EP_N^*$, where E is the nominal exchange rate. The real exchange rate is defined as $e = EP^*/P$.

When the US government levies a tariff $\tau > 0$ on the Japanese goods, the after-tax Japanese exports decrease

$$C_T^{*'} = \frac{C_T^*}{1+\tau} < C_T^*.$$

The Japanese producers thus can sell less units of goods. Since the labor market is competitive, the production sector already made zero profit prior to the US tariff. In response to the US tariff, the Japanese producers have no choice but to raise the price of their goods in order to pay off the wages. Hence, the after-tax Japanese aggregate price level becomes

$$P^{*'} = \Phi(\tau)P^* > P^*,$$

where $\Phi(\cdot)$ is a function of τ that should be larger than 1 but smaller than $(1+\tau)$. It is because the price of non-tradable goods will also increase in order to compensate the loss due to the reduction in sales. The after-tax real exchange rate becomes

$$e' = \frac{EP^{*'}}{P} = \frac{E[\Phi(\tau)P^*]}{P} = \Phi(\tau)e > e,$$

implying a yen appreciation and a dollar depreciation.

B Exogenous Risk-Bearing Capacity

Here we show the derivation of a simplified two-period model where the risk-bearing capacity is rather exogenous. The demand for dollar versus yen must be cleared in each period. We define the demand function of financiers Q_0 and the following market-clearing conditions for the dollar-yen at times $t = 0, 1$ as follows

$$e_0/(1+\tau^g) - \iota_0 + Q_0 = 0, \tag{B.1}$$

$$e_1 - \iota_1 - RQ_0 = 0, \tag{B.2}$$

where the time index of τ^g is suppressed in this two-period model.

Note that in (B.2), the export without the tariff term is at its frictionless level. There are two scenarios in reality that can justify this setup. First, the political party in the White House is expected to change, so the US trade policy will be reversed in the future. Second, the foreign trade partner (i.e. Japan in our example) levies their tariff on the US products as a retaliation response. Both scenarios result in a tariff reversal, implying that the economy will adjust towards the same direction as in the long-run frictionless equilibrium.

Proposition 4. *The equilibrium exchange rates follow*

$$e_0 = \frac{(1+\Gamma)\iota_0 + \iota_1}{1 + \frac{1+\Gamma}{1+\tau^g}}, \tag{B.3}$$

$$e_1 = \frac{\iota_0 + \left(1 + \frac{\Gamma}{1+\tau^g}\right)\iota_1}{1 + \frac{1+\Gamma}{1+\tau^g}}. \tag{B.4}$$

Since the world tariff only appears in the denominator of e_0 in (B.3), it is obvious to see that e_0 is increasing in τ^g, implying higher world tariff is associated with currency appreciation in the foreign countries.

Note that τ^g appears in both numerator and denominator of e_1 in (B.4), we take the partial derivative

$$\frac{\partial e_1}{\partial \tau^g} = \frac{(1+\Gamma)\iota_0 + \iota_1}{\left(1 + \frac{1+\Gamma}{1+\tau^g}\right)^2} > 0, \tag{B.5}$$

given $\Gamma, \iota_0, \iota_1 > 0$ by construction. To compare the magnitudes of currency appreciation across periods, we take partial derivative of exchange rate *return*, denoted as

$$\Delta e_1 = \frac{e_1 - e_0}{e_0} = \frac{\frac{\iota_1}{1+\tau^g} - \iota_0}{(1+\Gamma)\iota_0 + \iota_1}. \tag{B.6}$$

It is clear that the exchange change rate return is decreasing in the world tariff, implying that the currency appreciates, i.e. the dollar depreciates, more in the current period than in the future. It is not surprising as the impact of the world tariff wears off when the economy converges to the frictionless equilibrium. Nevertheless, we remark that the tariff's influence is long lasting even its appearance is only transitory.

C Notes

In Proposition 4, we can take partial derivatives of equilibrium exchange rates in (B.3) and (B.4) with respect to the current imports $\tilde{\iota}_0(\tau)$, which is decreasing in the world tariff, i.e. $\tilde{\iota}_0'(\tau) < 0$.[11]

$$\frac{\partial e_0}{\partial \tilde{\iota}_0(\tau)} = 1; \quad \frac{\partial e_1}{\partial \tilde{\iota}_0(\tau)} = 1/(1+\Gamma).$$

Since both derivatives are positive, it implies dollar *appreciations* ($e_0, e_1 \downarrow$) when tariffs are high ($\tau \uparrow$) in both periods. This is aligned with the findings in Jeanne and Son (2021)

[11] This negative relationship between the tariff and the imports (as part of the consumption) is found by Fender and Yip (2000) in a general equilibrium, two-country model. The authors argue that the negative impact of the rising tariff on the imposer's output is not only found in the short run but also in the steady state. Our model generates similar results on the equilibrium exchange rates.

and Matveev and Ruge-Murcia (2021). Nevertheless, the magnitude of appreciation is not identical given a fixed degree of rising tariffs. The dollar appreciates more in the short run than in the long run. As a result, the expected return on holding the dollar is negative. We remark an overshooting effect here in the presence of rising tariffs.

Alternatively, one could rewrite the model by focusing net exports and the dollar demand from the foreign households' perspective. The model ingredients are similar to the baseline, except for assuming $\iota_t^* = 1$ and $\xi_t^* \neq 1$ for $t = 0, 1$.

The market clearing conditions similar to (B.1) and (B.2) are

$$\tilde{\xi}_0^*(\tau) - e_0 - Q_0 = 0, \tag{C.7}$$

$$\xi_1^* - e_1 + Q_0 = 0. \tag{C.8}$$

The Japanese exports to the US are invoiced in the dollar, so we can simply express the above conditions in dollar terms.For simplicity, we also assume that the financial asset Q_0 is invoiced in the dollar. The signs assigned to Q_0 are the opposite to the market clearing conditions in the US as they are to clear the supply or demand from the US. Adding up (C.7) and (C.8), we obtain

$$\tilde{\xi}_0^*(\tau) + \xi_1^* = e_0 + e_1. \tag{C.9}$$

The financier's optimal condition remains the same as (5). The equilibrium exchange rates are

$$e_0 = \frac{(1+\Gamma)\tilde{\xi}_0^*(\tau) + \xi_1^*}{2+\Gamma}, \tag{C.10}$$

$$e_1 = \frac{\tilde{\xi}_0^*(\tau) + (1+\Gamma)\xi_1^*}{2+\Gamma}. \tag{C.11}$$

Similarly, the higher tariffs are associated with the dollar appreciations and the yen depreciation ($e_0, e_1 \downarrow$) since the net exports of the foreign countries are penalized ($\tilde{\xi}_0^*(\tau) \downarrow$). The expected return on exchange rate is

$$\Delta e_1 = -\frac{\Gamma}{2+\Gamma}(\tilde{\xi}_0^*(\tau) - \xi_1^*), \tag{C.12}$$

where the higher tariffs and thus the falling net exports ($\tilde{\xi}_0^*(\tau) \downarrow$) are associated with the dollar depreciation ($\Delta e_1 \uparrow$) and the yen appreciation. This is the same overshooting effect presented in the model of Jeanne and Son (2021). In another words, the magnitude of dollar appreciation due to the rising tariffs is larger currently at $t = 0$ than in the next period.

D Data Construction of Tariff Measures

In this section, we describe in details the constructions of tariff measures used in this paper. There are 4 measures considered in the analysis, including the import duties in the local currency, and in the USD, two types of average tariffs.

First, we collect "Customs and other import duties (current LCU)" from World Bank and denote as *Duties* henceforth. Second, we convert the duties into the unit of USD since it is not directly available from the World Bank. To ensure the accuracy of this conversion, we collect the GDP (current LCU and current US$) from the same database where both local currency and USD variables are available. We use the ratio of these two variables to convert the duties in local currency into the USD as the following

$$\text{Duties (US\$)} = \frac{\text{GDP (US\$)}}{\text{GDP (LCU)}} * \text{Duties (LCU)}.$$

Third, we follow Lohmann and O'Halloran (1994) to construct the *average* tariff, defined as the duties divided by the imports. For the imports data, we collect from the IMF's Direction of Trade Statistics for the following two reasons. The data is available at monthly frequency which matches our baseline analysis. In addition, we can extract the bilateral trade data. Since this paper is studying the exchange rates against US dollar, we collect the exports from the U.S. equivalent to the bilateral imports with the U.S. in the foreign countries. We then divide the *Duties (US$)* by the collected imports to obtain the average tariff.

Forth, we collect the tariff data "Tariff rate, applied, weighted mean, all products (%)" from the World Bank database. This is the World Bank staff estimates using various databases that contain the information of trade, the applied tariff schedules at product level. Weighted mean applied tariff is the average of effectively applied rates weighted by the product import shares corresponding to each partner country. This is different from the most favor nation rates or the bond rates which may not reflect the actual tariff variations. Nevertheless, when the effectively applied rate is unavailable, the most favored nation rate is used instead. As the duties data from World Bank, the weighted-average tariff rate is only available at annual frequency, and the European Economic Area (EEA) countries all have the same time series since 1988.

Finally, to combine the tariff measures of annual frequency in the analysis at monthly frequency, we use forward filling when there is missing values. We assume the annual observations occur at the end of the year, i.e. December. Hence, from the December of year t to the next November of year $t+1$, we fill the same value of tariff in year t.

Table A.1. Controlling for Local Political Cycles: Panel Regressions

This table presents estimates, for currency fixed effects panel regressions, of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US and/or a dummy variable FC that takes on the value of one (zero) during Centre-Left (Centre-Right) Political terms in major foreign countries like Canada, France, Germany, Italy, Japan, and the UK. Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). An US presidential cycle starts in November when the elections take place and ends four years after in October. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. The sample consists of monthly observations between October 1983 and October 2020 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
DP	−5.525**							−5.538**	−5.137*	−5.717**	−5.286**	−5.555**	−8.014**	−7.466**
	(2.599)							(2.601)	(2.633)	(2.585)	(2.604)	(2.584)	(2.838)	(3.294)
Canada		−1.308						−1.364						−2.525
		(2.712)						(2.698)						(3.063)
France			4.007						2.916					4.913
			(2.753)						(2.756)					(4.443)
Germany				−0.628						−1.428				2.980
				(2.565)						(2.512)				(4.681)
Italy					2.636						1.901			4.224
					(2.655)						(2.633)			(5.278)
Japan						−2.691						−2.758		1.190
						(2.713)						(2.715)		(3.810)
UK							2.032						7.087*	10.001*
							(3.771)						(4.166)	(5.049)
α	1.456	−0.437	−1.926	−0.816	−2.118	0.523	−1.466	2.187	0.696	2.265	0.633	3.166	1.457	−2.725
	(1.711)	(2.203)	(1.451)	(1.877)	(1.599)	(2.245)	(1.329)	(2.476)	(1.969)	(2.220)	(2.043)	(2.442)	(1.711)	(5.683)
R^2 (%)	1.599	1.165	1.291	1.145	1.234	1.239	1.174	1.323	1.374	1.325	1.344	1.401	1.624	1.885
N	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460	8,460

Table A.2. Controlling for US Business Cycle Fluctuations: Panel Regressions

This table presents panel regression estimates of nominal exchange rate returns regressed on a dummy variable DP that takes on the value of one (zero) during Democratic (Republican) Presidential terms in the US while controlling for the term spread (TSP), default spread (DSP), relative interest rate (RR), and the log dividend-price ratio (LDP). Exchange rates are defined as units of US dollars per unit of foreign currency such that a negative (positive) return denotes US dollar appreciation (depreciation). The control variables are demeaned and lagged relative to the exchange rate returns. A presidential cycle starts in November when the elections take place and ends four years after in October. Each specification includes a currency fixed effect. Standard errors (in parentheses) are clustered by currency and time (calendar month) dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively. Exchange rate returns are expressed in percentage per annum. The sample consists of monthly observations between October 1983 and October 2020 for a cross-section of 25 developed and emerging currencies. Exchange rates are from Datastream, the dividend-price ratio is from Robert Shiller's website, and the other data are from the FRED database.

	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
	Panel A: Controls lagged by one month					Panel B: Controls lagged by three months				
DP	−5.836**	−5.288*	−5.336**	−5.261**	−6.157**	−6.047**	−5.175*	−5.313**	−5.382**	−6.404**
	(2.615)	(2.709)	(2.635)	(2.634)	(2.779)	(2.593)	(2.661)	(2.633)	(2.611)	(2.678)
TSP	1.176				1.124	1.521				1.379
	(1.168)				(1.250)	(1.176)				(1.249)
DSP		−5.898			−6.487		−8.060			−10.101
		(5.944)			(6.274)		(5.281)			(5.551)
RR			−0.770		2.243			−1.073		3.285
			(2.060)		(2.090)			(2.015)		(1.986)
LDP				2.726	0.088				1.716	−3.735
				(4.170)	(4.141)				(4.053)	(4.223)
α	1.685	1.436	1.406	1.452	1.926	1.794	1.338	1.393	1.475	1.786
	(1.761)	(1.769)	(1.691)	(1.713)	(1.848)	(1.765)	(1.773)	(1.706)	(1.715)	(1.861)
R^2 (%)	1.380	1.599	1.288	1.313	1.737	1.449	1.879	1.303	1.289	2.154
N	9,348	8,960	9,348	9,348	8,960	9,348	8,924	9,348	9,348	8,924

(continued)

Table A.2. Controlling for US Business Cycle Fluctuations: Panel Regressions *(continued)*

	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
	Panel C: Controls lagged by six months					Panel D: Controls lagged by one year				
DP	−6.083**	−5.209**	−5.264**	−5.382**	−5.879**	−6.590**	−5.198**	−5.660**	−5.478**	−6.297**
	(2.604)	(2.605)	(2.596)	(2.588)	(2.638)	(2.600)	(2.648)	(2.587)	(2.588)	(2.707)
TSP	1.283				0.858	2.079				1.831
	(1.195)				(1.341)	(1.136)				(1.327)
DSP		−9.529			−11.142		−5.295			−4.858
		(3.621)			(3.903)		(3.349)			(3.772)
RR			−2.264		2.204			−1.596		−0.116
			(2.489)		(2.721)			(2.237)		(2.718)
LDP				2.251	−3.333				1.458	−3.439
				(3.804)	(4.227)				(3.563)	(4.320)
α	1.794	1.328	1.386	1.491	1.481	2.035	1.154	1.590	1.511	1.611
	(1.788)	(1.769)	(1.704)	(1.722)	(1.891)	(1.748)	(1.783)	(1.742)	(1.726)	(1.950)
R^2 (%)	1.395	2.121	1.409	1.302	2.236	1.585	1.443	1.343	1.286	1.643
N	9,348	8,870	9,348	9,348	8,870	9,348	8,762	9,348	9,348	8,762

Table A.3. Implied Volatility and Trade Events: Median Statistics

This table presents pool and panel regression estimates of foreign exchange options' implied volatility differences centred around selected trade events. For each column, the reported statistics are the median of 33 regressions in which one trade event is removed at a time. The implied volatility differences are based on a window of one week (i.e., three days before and another three days after the trade event) for maturities ranging between one week and two years. 10δ Put (10δ Call) denotes the implied volatility of a deep out-of-the-money option that gives the right to sell (buy) a unit of foreign currency in exchange of US dollars whereas 25δ Put (25δ Call) refers to the implied volatility of an out-of-the-money option that gives the right to sell (buy) a unit of foreign currency in exchange of US dollars. ATM indicates the implied volatility of a delta-neutral straddle, commonly referred to as at-the-money. The set of independent variables includes country size and distance where the size of each country is rescaled by the total GDP of all countries in our sample, and the distance is expressed in thousand kilometres between the US capital and the Foreign country's capital. Standard errors (in parentheses) are clustered by currency and maturity dimension. The superscripts *, **, and *** indicate statistical significance at 10%, 5%, and 1% respectively, according to the median of t-statistics. The sample consists of monthly observations between January 1996 and May 2020 for a cross-section of 19 developed and emerging currencies. Foreign exchange options' implied volatility data are from JP Morgan and Bloomberg. GDP data are from the World Economic Outlook Database and distance data are hand collected.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	10δ Put		25δ Put		ATM		25δ Call		10δ Call	
Country Size	0.034	0.034	0.034	0.034	0.035	0.035	0.036	0.036	0.036	0.036
	(0.038)	(0.048)	(0.038)	(0.048)	(0.038)	(0.048)	(0.037)	(0.048)	(0.037)	(0.048)
Distance	−0.002	−0.002	−0.002	−0.002	−0.002	−0.002	−0.003	−0.003	−0.003	−0.003
	(0.022)	(0.023)	(0.022)	(0.023)	(0.022)	(0.023)	(0.022)	(0.023)	(0.022)	(0.023)
α	0.861***	0.861***	0.768***	0.768***	0.687***	0.687***	0.618***	0.618***	0.591***	0.591***
	(0.244)	(0.190)	(0.230)	(0.172)	(0.215)	(0.153)	(0.198)	(0.139)	(0.189)	(0.132)
R^2 (%)	0.008	1.830	0.020	2.404	0.046	3.036	0.086	3.182	0.083	2.946
N	4133	4133	4135	4135	4135	4135	4135	4135	4135	4135
Maturity FE	N	Y	N	Y	N	Y	N	Y	N	Y