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Exchange rates and political uncertainty: the Brexit case

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Abstract

This paper studies the impact of political risk on exchange rates. We focus on the Brexit Referendum as it provides a natural experiment where both exchange rate expectations and a time-varying political risk factor can be measured directly. We build a portfolio model that relates changes in the Leave probability to changes of the British pound's market price, both via expectations and via a political risk factor. We estimate the model for multilateral and bilateral British pound exchange rates. We find that the Leave probability predicts a depreciation of the pound, consistent with the outcome post-referendum, and that the time-varying political risk affects exchange rates independently.

1 | INTRODUCTION

Politics has long been recognized as a major determinant of the international price of currencies. However, political risk is notoriously difficult to quantify. To overcome this hurdle, recent work proposed using indices of political uncertainty that aggregate over multiple sources of information to generate widely comparable and relatively long time series (e.g. Baker *et al.* 2016). To complement this literature, and to dig into the mechanism by which political risk is priced, this paper takes a different route: it focuses on a major recent political event—the Brexit Referendum in Great Britain—for which currencies' response to time-varying political risk can be measured directly. Our findings confirm that a political risk premium plays a crucial role in exchange rate determination.

The Brexit Referendum shook the European and international political scene in 2016: for the first time, a European Union member country, the UK, voted for leaving the Union. The debate in the run up to the referendum focused on mainly international issues, trade and immigration, which motivates our interest in its implications for exchange rates. With few exceptions, economists agreed that Brexit would have negative consequences on the British economy (e.g. Sampson *et al.* 2016)[1] and expressed concerns that the City of London would lose its role as

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the main market for euro denominated assets.[2] Indeed, with the victory of the 'Leave' camp, the British pound (BP) depreciated overnight by about 7% against the euro and other main currencies.

What makes the Brexit Referendum an interesting 'natural experiment' where political risk can be measured directly is that the event has been preceded by an exceptionally liquid betting market.[3] Bookmakers such as Betfair and PredictIt provided online platforms on which individuals could bet on the likely outcome of the referendum in real time. From the bookmakers' Brexit odds, we construct a daily series of the Leave probability. Our first question is: *does the evolution in the odds-implied Leave probability relate to movements in the price of the BP relative to the other currencies?*

To answer, we start by writing down a simple model where risk-neutral investors form expectations about post-referendum exchange rates. In the absence of arbitrage opportunities and transaction costs, the interest rate on a domestic currency denominated asset should exceed the foreign one when the domestic currency is expected to depreciate. We write this uncovered interest parity (UIP) condition allowing the expected depreciation to depend on the (time-varying) probability of Leave. We then take this model to the data and find that market participants expected the BP to depreciate against all major currencies upon a victory for Leave. Specifically, we predict a depreciation of the BP against a basket of major currencies of approximately 15% in case of a Leave victory.[4]

While the sign of the relation between the Leave probability and a depreciation of the BP is consistent with what happened in the currency markets after the referendum, our first model tends to *overestimate* considerably the actual depreciation, by about 7–8 percentage points. To shed some light on this, we take advantage of an important—yet thus far overlooked—property of odds-implied events probabilities: the fact that they contain information not only on the expected effect of an event on asset prices (first moment), but also on the time-varying *political risk* (second moment). Intuitively, the closer the event probability is to one-half, the larger is the political uncertainty faced by market participants, the relationship being non-linear. Thus when the outcome of the referendum becomes known, political uncertainty actually falls, and this effect could contribute to a currency *appreciation*, which partially offsets the expectations effect.

Indeed, this is what we find when we extend our model by allowing the marginal investor in the currency market to be risk averse. We derive a closed-form solution for the time-varying (political) risk premium, which turns out to be a non-linear function of the Leave probability. Our second question is: *do currency markets price our (model-based) measure of political risk premium, and if so, can this explain the exchange rate under-reaction?* We find that our measure of political risk is positively and highly significantly associated with movements in all exchange rates considered—with the exception of the BP versus Japanese yen—and that the reduction of uncertainty in the aftermath of the referendum accounts for an appreciation of the BP by about 8%. This confirms that two channels are at play. First, expectations (first moment): when the Leave scenario is more likely, investors expect a depreciation after the referendum, so the exchange rate weakens immediately.[5] Second, political risk (second moment): when the probability of Leave gets closer to 50%, the risk of investing in the BP reaches a maximum, and investors re-allocate their portfolios away from the BP into other currencies.

Interestingly, the effect of political risk in our paper is very different from that of the typical time-varying risk aversion channel, according to which risk aversion increases when a bad state of the world occurs (see, for example, Guiso *et al.* 2018). Provided that market participants view Brexit as having negative effects on the BP and on the UK economy, we would expect their risk aversion to increase, not decrease, in the aftermath of the referendum. In other words, the time-varying risk aversion argument predicts an over-reaction to a negative shock, not an under-reaction. In contrast, our political risk argument is consistent with the data: following the referendum, the uncertainty about its outcome actually falls (while deep parameters such as risk aversion stay put). Our story is consistent with what happened following Boris Johnson's electoral victory in December 2019 (see Manasse *et al.* 2020), when the BP appreciated by 2%. In our interpretation, Johnson received a clear mandate to end the uncertainty related to the deadlock of negotiations with the EU and to proceed without further ado to Brexit.

The last question that we ask is: *can we improve on the standard single-equation exchange rate models by considering a simultaneous, multi-currency portfolio setting?* This approach has the advantage that as long as investors choose the currency composition of their portfolios by considering interest yields and the covariance structure of currencies, their risk attitude should be the same across all currencies. As a consequence, all bilateral exchange rates should be affected equally by a change in the political risk premium associated with Brexit. When we estimate a dynamic seemingly unrelated regressions (SUR) system of equations for multiple currencies, we find that our cross-equation restriction is not rejected by the data. Relative to single-equation regressions, SUR estimates are more precise, and the UIP parameter is closer to the theoretical value of one for both the BP exchange rate with the euro and with the US dollar.

The paper unfolds as follows. Section II reviews the relevant literature. Section III presents the theoretical model, first under risk neutrality and then allowing for risk aversion and multiple currencies. Section IV presents the data. Section V discusses the empirical strategy and presents the results, as well as a number of robustness checks. Section VI concludes.

2 | LITERATURE REVIEW

Our paper contributes to several strands of the literature. First, we complement the recent work that measures political uncertainty via aggregate indices (e.g. Baker *et al.* 2016; Pastor and Veronesi 2012; Brogaard and Detzel 2015; Fernández-Villaverde *et al.* 2015; Kelly *et al.* 2016), by constructing an alternative political-risk measure that both has a theoretical interpretation and can be observed directly from bookmakers' odds—a market price that conveys the 'wisdom of the crowd'. Related work on micro-measuring the price of political uncertainty has focused on either tax policies (Sialm 2009; Croce *et al.* 2012) or equity premia (Santa-Clara and Valkanov 2003; Belo *et al.* 2013; Bittlingmayer 1998; Voth 2003; Boutchkova *et al.* 2012).

In addition, our results relate to the vast literature on exchange rates. While recent work analysed the pricing of macroeconomic uncertainty in currency markets (e.g. Ismailov and Rossi 2018; Rossi and Sekhposyan 2015), the role of political uncertainty has been largely overlooked. To our knowledge, the only exception is Bachman (1992), who studies the impact of political news (elections) on the time-varying risk premium in foreign exchange markets. Before an election, there is uncertainty about whether the new government will implement policy (a tax on domestic assets) that affects the domestic interest rate. This uncertainty is resolved after the elections. Thus the parameters of an exchange rate equation should be unstable: those estimated in the sample before the elections should differ significantly from those in the post-election period. Unlike Bachman (1992), we do not limit ourselves to testing for structural breaks, but we estimate directly the effects of time-varying political uncertainty.

Our paper is also related to the vast literature on exchange rates and the UIP anomaly, starting from the seminal contributions of Fama (1984), Hodrick (1987) and Froot and Frankel (1989), and which includes recent contributions by Burnside *et al.* (2006), Chinn and Quayyum (2012), Engel (2014) and Ismailov and Rossi (2018). Instead of positing a time-varying risk premium solution (as in Verdelhan 2010; Lustig *et al.* 2011; Bansal and Shaliastovich 2013; Farhi and Gabaix 2015) or deviating from standard, rational expectations (as in Gourinchas and Tornell 2004; Burnside *et al.* 2011; Ilut 2012), we focus on a specific 'natural experiment' in which we can measure them directly from bookmakers' odds. The more recent theoretical and empirical literature on exchange rates develops from the paper by Evans and Lyons (2002), which introduces micro-structure finance determinants of exchange rates by looking at inter-dealers currency order flows. In their model, these flows, defined as the difference between the numbers of buyers' and sellers' initiated orders, convey information on exchange rates. This is because specialized dealers intermediate the private net demands for currency. In our paper, we use directly micro-price data from the parallel market of bets to construct the relevant probabilities underlying exchange rate expectations in the wake of the Brexit referendum. More recently, Kalemli-Özcan and Varela (2021) find that political uncertainty, as captured by the Economic Policy Uncertainty (EPU) measure of Baker *et al.* (2016), is an important determinant of systematic deviations from the UIP across countries, which is consistent with our findings for the Brexit experiment.

In addition, our paper is related to Itskhoki and Mukhin (2021). They develop a general equilibrium model to explain the most intriguing exchange rate 'puzzles' in the international macro-finance literature.[6] Building on De Long *et al.* (1990), Jeanne and Rose (2002), and Gabaix and Maggiori (2015), they combine noise traders and risk-averse arbitrageurs, and obtain a modified interest parity condition that features an endogenous risk premium. This depends on the arbitrageurs' risk aversion parameter and on the exchange rate volatility (equation (16) in their paper). In our paper, which takes a partial equilibrium approach to focus on political risk, we solve for the optimal portfolio under risk aversion, and obtain a very similar condition; see our equation (11), where the time-varying risk premium is entirely a reflection of political risk, for given risk aversion.

Finally, our paper is related to the literature that discusses possible econometric reasons for the failure of the uncovered parity condition. This includes the contributions of Meese and Rogoff (1983a,b), Baillie and Bollerslev (2000), Cheung *et al.* (2005), Chinn and Meredith (2005), Rossi (2006), Alquist and Chinn (2008), Brunnermeier *et al.* (2008), Clarida *et al.* (2009), Chinn and Quayyum (2012), Menkhoff *et al.* (2012) and Chen and Tsang (2013). When we extend the model to multiple currencies and estimate a SUR system of equations, we find that, relative to the single-equation model, the SUR estimates are more precise and the interest-rate parameter is closer to the *a priori* correct value for both the EU and the USA.

3 | THE MODEL

In this section, we lay out our baseline model. We start by assuming risk neutrality, under the simplifying assumption that there is binary uncertainty over the outcome of either Leave or Remain votes. Later, we introduce risk aversion to derive an observable measure of the time-varying political risk premium. Finally, we consider a portfolio model with many currencies, where the parameters of different exchange rate regressions are not independent and satisfy a cross-equation restriction that we test on the data.

3.1 | Risk neutrality, two currencies

This subsection develops the simplest framework where we can relate exchange rate movements to the resolution of uncertainty around an event. Denote by i (i^*) the UK (foreign) interest rate on assets that are issued before, and mature after, the date of the referendum.[7] The (natural logarithm of the) nominal exchange rates before and after the resolution of uncertainty on the referendum outcome are denoted by e and e', respectively, and are expressed as units of the domestic currency (BP) for one unit of foreign currency. The uncovered parity condition states that the difference between the interest rates of a domestic and foreign currency denominated asset must equal expected depreciation, which we write as

$$e = i^* - i + \mathbb{E}(e'). \tag{1}$$

For now, we assume that a referendum will certainly take place. Given that a referendum has only two possible outcomes, $V = \{\text{Leave } (L), \text{Remain } (R)\}$, and assuming that the referendum result is the only relevant uncertain event for the exchange rate around the referendum day, we can write the future expected exchange rate as

$$\mathbb{E}(e') = \pi \ \mathbb{E}(e'|L) + (1-\pi) \ \mathbb{E}(e'|R), \tag{2}$$

where $\mathbb{E}(e'|V)$ is the expected future exchange rate conditional on outcome V, and π is the probability of Leave. Substituting the previous definition into equation (1) yields

$$e = i^* - i + \pi \ \left[\mathbb{E}(e'|L) - \mathbb{E}(e'|R)\right] + \mathbb{E}(e'|R). \tag{3}$$

Assume that the probability of a Leave outcome can be observed—for instance, from betting odds posted by risk-neutral wagerers. Then a linear regression of the exchange rate on the foreign–domestic spread and on the odds probability yields

$$e = \theta(i^* - i) + \alpha + \beta\pi, \tag{4}$$

where equation (3) implies

$$\alpha = \mathbb{E}(e'|R), \tag{5}$$

$$\beta = \mathbb{E}(e'|L) - \mathbb{E}(e'|R), \tag{6}$$

$$\theta = 1. \tag{7}$$

The intercept α equals the market's expectation of the future exchange rate conditional on Remain, and the sum of the estimates of the intercept and slope coefficients $\alpha + \beta = \mathbb{E}(e'|L)$ equals the expected exchange rate conditional on Leave. A positive (negative) estimate of the coefficient β implies that the exchange rate is expected to depreciate (appreciate) after the referendum if Leave wins. The probability of Leave (π) takes value 1 or 0 after the referendum result becomes known, as this source of uncertainty is resolved.

3.2 | Risk aversion, two currencies

Now we allow for risk aversion and derive a measure of the political risk premium. Consider a representative foreign (i.e. non-UK) investor that chooses between investing in its own currency denominated asset, which is safe, or in a BP denominated asset, which is risky because of exchange rate fluctuations due to political uncertainty. The portfolio choice is made before the resolution of uncertainty that occurs with the vote of the referendum, as the assets mature after the vote. Let the portfolio share that is invested in BP denominated assets be ω. For simplicity, we consider standard mean–variance preferences, so that the investor chooses ω in order to maximize

$$U(\omega) = (1-\omega)i^* + \omega \left[i - (\mathbb{E}(e') - e)\right] - \frac{1}{2} r\omega^2\sigma^2, \tag{8}$$

where $r/2 \geq 0$ is the coefficient of absolute risk aversion, and σ^2 is the portfolio variance. In equation (8), the first term on the the right-hand side represents the return on the risk-free non-UK (safe) investment, i^*, while the second term represents the expected return on the UK (risky) investment, given by the sum of the interest rate i and the expected appreciation of the BP. Each term is weighted by its respective portfolio share. The last term in the utility function is the product of the risk aversion parameter $r/2$ multiplied by the (squared) share of UK assets times the risk on the portfolio return, σ^2.[8] The portfolio variance can be written as

$$\begin{aligned}\sigma^2 &= \pi\ \mathbb{E}(e'|L)^2 + (1-\pi)\ \mathbb{E}(e'|R)^2 - \left[\pi\ \mathbb{E}(e'|L) + (1-\pi)\ \mathbb{E}(e'|R)\right]^2 \\ &= \pi(1-\pi)\left[\mathbb{E}(e'|L) - \mathbb{E}(e'|R)\right]^2 \\ &= \pi(1-\pi)\beta^2. \end{aligned} \tag{9}$$

Equation (9) shows that the portfolio variance coincides with the volatility of the exchange rate. This volatility increases with the difference between the expected exchange rates conditional on the two possible outcomes β (see equation (6)), and reaches a maximum when the Leave/Remain probabilities are the same, or $\pi = \frac{1}{2}$. From the first-order condition, the optimal portfolio share of UK denominated investments reads

$$\omega = \frac{i - [\mathbb{E}(e') - e] - i^*}{r\sigma^2}. \tag{10}$$

Intuitively, the share in the (risky) BP denominated assets is an increasing function of the difference between the UK and the foreign expected yields. With risk-averse investors (i.e. $r > 0$), the share of BP denominated assets is decreasing in the portfolio volatility σ^2 and in the degree of risk aversion. We can solve for the exchange rate e by assuming that the asset market clears, so that the demand for foreign assets (ω) equals supply (s), which we take as *exogenous*.[9] Substituting $\omega = s$ in equation (10) and solving for the exchange rate, we obtain a modified uncovered parity condition

$$e = \theta(i^* - i) + \alpha + \beta\pi + \gamma\pi(1-\pi), \tag{11}$$

where

$$\alpha = \mathbb{E}(e'|R) > 0, \quad \beta = \mathbb{E}(e'|L) - \mathbb{E}(e'|R), \quad \gamma = r\beta^2 s \geq 0, \quad \theta = 1. \tag{12}$$

This expression generalizes equation (4) for the case of risk aversion, with the addition of a time-varying risk premium given by $r\sigma^2 s = r\pi(1-\pi)\beta^2 s$. This premium is increasing in the equilibrium portfolio share of BP (i.e. s), in the coefficient of absolute risk aversion ($r/2$), and in the volatility of the portfolio return (σ^2). The risk premium will be time-varying as long as the probability of Leave (π) changes over time.

Note that a *marginal* increase in the Leave probability (from below $\frac{1}{2}$) now has two separate effects on the current exchange rate. On the one hand, it implies a rise in the future expected exchange rate—provided that the investors forecast a weaker BP in case of Brexit, that is, $\beta > 0$. On the other hand, a larger Leave probability raises the political uncertainty, leading risk-averse investors to reduce their portfolio share of BP and further depreciating the exchange rate. Thus a small increase in the Leave probability π raises the current exchange rate e by more than it raises the future expected exchange rate $\mathbb{E}(e')$, and this *overshooting* implies an expected *appreciation* of the currency. As a result, a marginal change in the Leave probability induces a *negative correlation* between the risk premium and the expected *depreciation* rate, as required by Fama (1984) to explain the UIP puzzle. Note, however, that a *discrete* increase of the Leave probability, from below $\frac{1}{2}$ to 1, following the resolution of uncertainty after a Leave victory, may actually imply an opposite effect. In particular, the BP may appreciate, provided that investors are sufficiently risk averse, they hold a large share of BP denominated assets, and the percentage difference in conditional exchange rate expectations prior to the resolution of uncertainty is large enough (i.e. $\beta < \gamma$ if $r\beta\omega > 1$).

In the Appendix, we generalize this analysis to the case in which some residual uncertainty—for instance between a 'hard' and a 'soft' Brexit—persists after the outcome of the referendum becomes known. In this case, after a Leave outcome, the exchange rate volatility can either decrease (as happens in the current model) or increase, as the vote may open the way to an even more turbulent scenario ahead (a 'hard' one).[10]

3.3 | Referendum uncertainty, two currencies

In our framework, the parameters of the exchange rate equation (α, β, γ) depend on 'deep' parameters that reflect agents' information and preferences, and which may be subject to change when crucial Brexit-related news arrives. We dig further into this issue by considering an alternative reason for parameter instability: the resolution of uncertainty on the referendum occurrence, in addition to its eventual outcome. To this end, we extend the model by taking a step backwards in time when, following the UK negotiations with the EU, investors were uncertain on whether a referendum on Britain's membership of the EU would ever be called. This extension tries to rationalize what happened when the British Parliament promulgated the so-called 'Referendum Act' in mid-December 2015, which established the constitutionality of a referendum on Great Britain's exit from the EU, and made sure that a referendum would take place. As we will discuss later, three things happened in the following weeks: currency markets reacted with a sharp depreciation of the BP; the different bookmakers' odds on the outcome of the vote, which until then showed little correlation among themselves, started to converge and co-move; and finally, their correlation with the exchange rate increased sharply. Let q denote the probability that a referendum will take place, so that with probability $1-q$ the referendum will not take place, in which case the exchange rate is expected to be $\mathbb{E}(e'|\emptyset)$. In this case, the exchange rate expected when the investor does not know whether a referendum will take place is

$$\mathbb{E}(e') = (1-q)\ \mathbb{E}(e'|\emptyset) + q\left[\pi\ \mathbb{E}(e'|L) + (1-\pi)\ \mathbb{E}(e'|R)\right],$$

which is a linear combination of the expected exchange rate conditional on no referendum, the first term, and the expected exchange rate conditional on the referendum taking place. It is easy to show that for the case of risk-neutral investors, the exchange rate equation is modified as

$$e = \theta(i^* - i) + \alpha + q\delta + q\beta\pi,$$

where β and θ are defined as before, now $\alpha = \mathbb{E}(e'|\emptyset)$, and $\delta = \mathbb{E}(e'|R) - \mathbb{E}(e'|\emptyset)$. In particular, if the market anticipates that the exchange rate prevailing after a Remain vote will be the same as the one that would have prevailed had the referendum not occurred, $\delta = 0$, then the equation simplifies to

$$e = \theta(i^* - i) + \alpha + q\beta\pi. \tag{13}$$

Comparing equations (13) and (4), we can see that the coefficient of the Leave probability π should make a discrete jump from $q\beta$ to β at the moment when the Brexit Referendum is announced. Intuitively, at this moment, the effect of the odds on the expected exchange rate will increase. A similar intuition holds for the portfolio model with risk aversion. In this case, it is easy to show that introducing uncertainty on the actual occurrence of the referendum modifies the exchange rate equation to

$$e = \theta(i^* - i) + \alpha + q\delta + q\beta\pi + \gamma q\pi(1-\pi), \tag{14}$$

where the parameters β, γ and θ are the same as in (12), $\alpha = \mathbb{E}(e'|\emptyset)$, and $\delta = \mathbb{E}(e'|R) - \mathbb{E}(e'|\emptyset)$. This equation shows that also the coefficient of the risk premium $\pi(1-\pi)$ should exhibit a discrete upward jump at the time of the referendum announcement. The portfolio variance now reads $\sigma^2 = q\pi(1-\pi)\beta^2$.

3.4 | Risk aversion, many currencies

Our portfolio model generalizes immediately to the case of many currencies, where the (non-UK) investor can chose between $N+1$ assets, one denominated in his/her own currency, and N denominated in 'risky' foreign currencies, in the sense that their price in terms of the domestic currency will be affected, perhaps differently, by the referendum outcome. Their yields are denoted by i_j, $j = 1, \ldots, N$. The portfolio weights are denoted by ω_j. For simplicity, we consider the case where the referendum is known to take place for sure. Let e_j denote the price of currency j in terms of the investor's currency. Investing in a foreign currency j entails a political risk due to the referendum. We need to replace the previous objective function (8) with the expression

$$U = \left(1 - \sum_j \omega_j\right) i^* + \sum_j \omega_j \left(i_j - \left[\pi \, \mathbb{E}(e'_j|L) + (1-\pi)\, \mathbb{E}(e'_j|R) - e_j \right] \right)$$
$$- \frac{r}{2} \left(\sum_j \omega_j^2 \sigma_j^2 + 2 \sum_i \sum_{j>i} \omega_i \omega_j \sigma_{ij} \right),$$

where the portfolio variance now depends on the weighted sum of all bilateral variances and covariances σ_{ij}. Using the fact that

$$\mathbb{E}(e'_i e'_j) = \pi \, \mathbb{E}(e'_i|L)\, \mathbb{E}(e'_j|L) + (1-\pi)\, \mathbb{E}(e'_i|R)\, \mathbb{E}(e'_j|R),$$

we can write the covariance between two currencies i and j as

$$\begin{aligned}
\sigma_{ij} &= \mathbb{E}(e'_i e'_j) - \mathbb{E}(e'_i)\, \mathbb{E}(e'_j) \\
&= \pi(1-\pi) \left[\mathbb{E}(e'_i|L) - \mathbb{E}(e'_i|R) \right] \left[\mathbb{E}(e'_j|L) - \mathbb{E}(e'_j|R) \right] \\
&= \pi(1-\pi)\beta_i \beta_j.
\end{aligned}$$

This expression shows that the covariance between any pair of currencies is equal to the product of the expected 'jumps' β of the two currencies, multiplied by the uncertainty over the referendum outcome $\pi(1-\pi)$. This extends equation (9) to the case of many currencies. Proceeding as before, we can calculate the optimal portfolio shares ω_j, and derive the equilibrium price equations, for $j = 1, \ldots, N$, satisfying $\omega_j = s_j$. We obtain

$$e_j = i^* - i_j + \mathbb{E}\left(e'_j|R\right) + \pi \left[\mathbb{E}\left(e'_j|L\right) - \mathbb{E}\left(e'_j|R\right) \right] + r \sum_j \sum_{k>j} s_j s_k \sigma_{jk},$$

which we can rewrite in compact form as

$$e_j = \theta(i^* - i_j) + \alpha_j + \beta_j \pi + \tilde{\gamma} \pi (1-\pi), \tag{15}$$

where for $j = 1, \ldots, N$,

$$\alpha_j = \mathbb{E}\left(e'_j|R\right), \quad \beta_j = \mathbb{E}\left(e'_j|L\right) - \mathbb{E}\left(e'_j|R\right), \quad \tilde{\gamma} = r \sum_j \sum_{k>j} s_j s_k \beta_j \beta_k, \quad \theta = 1.$$

Expression (15) represents a system of N non-linear equations. These equations exhibit a cross-equation restriction on the parameters, which comes from the fact that, unlike the intercept α_j and the slope β_j, which are country-specific, the coefficient of the volatility term, $\tilde{\gamma}$, depends on the expected 'jumps' of *all* the exchange rates in the portfolio and on the risk aversion parameter, and should therefore be identical across currencies.

In this model, the coefficient of the political risk premium term, $\tilde{\gamma}$, in principle can be negative, so that a foreign currency, such as the BP, could 'appreciate' when political risk increases. This would apply if this currency were expected, say, to appreciate relative to the domestic currency while some other currency k were expected to depreciate, $\beta_j < 0$, $\beta_k > 0$, so that investing in the foreign currency may entail a diversification benefit that would reduce exposure to political risk. This extension suggests that imposing the restriction that the volatility coefficient should be the same across currencies and estimating a system of equation could improve the efficiency of the estimates.

4 | THE DATA

We collect daily data from 27 May 2015 to 23 June 2017 on exchange rates, interest rates and bookmakers' odds, as well as other measures of political and economic uncertainty that have been considered by previous research. We study the behaviour of the BP vis-à-vis the currencies of its major trading partners: the euro (EUR), the US dollar (USA), the Japanese yen (JAP), the Swiss franc (CHE), the Canadian dollar (CAN), the Danish krone (DAN), the Swedish krone (SWE), the Norwegian krone (NOR), the Australian dollar (AUS) and the New Zealand dollar (NZL). To this aim, we consider two specifications of a multilateral nominal exchange rate of the BP against a basket of these ten currencies, as well as each bilateral exchange rate.

Our first multilateral exchange rate and the corresponding interest rate differential are constructed as weighted averages of the individual country-specific exchange rates and interest rate differentials, respectively, using trade-based weights. In particular, the weight of each country is calculated as the ratio of the BP-valued bilateral trade (the sum of exports and imports) of the UK with that country and the total BP-valued trade of the UK with the ten countries altogether, measured in 2015.[11] We also construct an alternative measure of multilateral exchange rate based on financial weights. Specifically, for a generic country i, the financial weight is given by the ratio of the financial position (the sum of assets and liabilities) of the UK towards country i, relative to the total financial position of the UK against all the other ten countries, measured as of December 2015.

Table 1 reports the two sets of weights in columns (2) and (4), and the values of the bilateral financial position and bilateral trade in 2015 in columns (1) and (3), respectively. In both baskets, the euro and the US dollar (USD) play a dominant role. The trade weights reflect the predominant share of the euro area in UK trade (64%), with a relatively minor role for the USD (17%) and a non-negligible role for the Swiss franc (6%), while financial weights assign to the euro and the USD 47% and 36%, respectively, and about 6% to the yen.

The data source for exchange rates and interest rates is Datastream. We initially tried a variety of interest rates, inter-bank rates and treasury bill rates, of different maturities. We finally settled on the 3-month LIBOR inter-bank rate, as this is standard in the literature (see Ismailov and Rossi 2018) and also because the main results remained the same. The source for international financial positions is the IMF Coordinated Portfolio Investment Survey. International trade data are from the IMF Direction of Trade Statistics.

Our Leave probability measure is constructed using real time data on odds[12] provided by two betting companies: Betfair and PredictIt. For either provider, we take the daily average probability of Leave derived from the corresponding odds. The probability variable π is the average of the two series, but we also experiment with different weighting schemes. We chose to use this type of data as prediction markets are likely to reflect the 'wisdom of the crowd'. In particular, unlike survey data, they are immune from misreporting as investors 'put their money where their mouth is'. In the words of Arrow *et al.* (2008):

> because information is often widely dispersed among economic actors, it is highly desirable to find a mechanism to collect and aggregate that information. Free

TABLE 1 Financial and trade weights of the UK's economic partners in 2015.

	Financial weights		Trade weights	
	Billions US$	%	Billions US$	%
Country	(1)	(2)	(3)	(4)
EUR	2928.21	47.15	472.46	64.30
USA	2207.69	35.55	124.58	16.96
JAP	384.11	6.18	15.99	2.18
CHE	164.15	2.64	41.99	5.72
CAN	113.99	1.84	19.14	2.61
DAN	49.81	0.80	8.76	1.19
SWE	95.09	1.53	17.15	2.33
NOR	115.98	1.87	23.80	3.24
AUS	140.31	2.26	8.64	1.18
NZL	11.51	0.19	2.21	0.30

Notes: The table reports the weights used to construct the multilateral exchange rates for the BP against the basket of ten currencies considered in the paper. For a generic country *i*, the financial weight is given by the ratio of the financial position (the sum of assets and liabilities) of the UK towards country *i*, relative to the total financial position of the UK against all ten countries, measured as of December 2015. The trade weight is calculated as the ratio of the bilateral trade (the sum of exports and imports) between the UK and country *i*, relative to the value of UK trade with the ten countries, measured in 2015.
The source for international financial positions is the IMF Coordinated Portfolio Investment Survey. The source for international trade data is the IMF Direction of Trade Statistics.

> markets usually manage this process well because almost anyone can participate, and the potential for profit (and loss) creates strong incentives to search for better information. (Arrow *et al.* 2008)

Prediction markets are used to manage risks—such as flu outbreaks and environmental disasters—by public entities (e.g. US Department of Defense) as well as firms (e.g. General Electric, Google, Hewlett-Packard, IBM, Microsoft).

We obtained data from two of the major online betting markets that were active in 2016. Betfair is a British online gambling company headquartered in Hammersmith (West London) and Clonskeagh (Dublin). It claims to have over 4 million customers (1.1 million active customers) and a turnover in excess of £50 million per week. As of April 2013, the company employed 1800 people. On its betting website, Betfair listed two Brexit-related contracts: the first paid out £1 in the event of a Leave victory; the other paid £1 in the event of Remain. Betfair supplied us with the odds implicit in the contract price, which are observed from 27 May 2015 to 24 June 2016 at different time intervals (often of 1 second), each weekday, for a total of 143,289 observations. This data source was used in recent papers (e.g. Auld and Linton 2019). We complement this with a second source for betting odds, the New Zealand based company PredictIt, which launched a market on the Brexit vote on 3 November 2014, and caters for mainly US-based investors.[13]

Figure 1 compares the two Leave probability series derived from the Betfair odds (solid blue line) and PredictIt ones (dashed green line). The figure reveals that the Leave probability measures were quite noisy in 2015: until December 2015, the two series exhibit a negative correlation (−0.35), and the standard deviation of their difference is 9.3%. However, starting in January 2016, the odds appear to behave similarly: for the period 1 January to 22 June, the standard deviation of the difference between the series falls to 4.2%, and the correlation coefficient rises to 0.51.

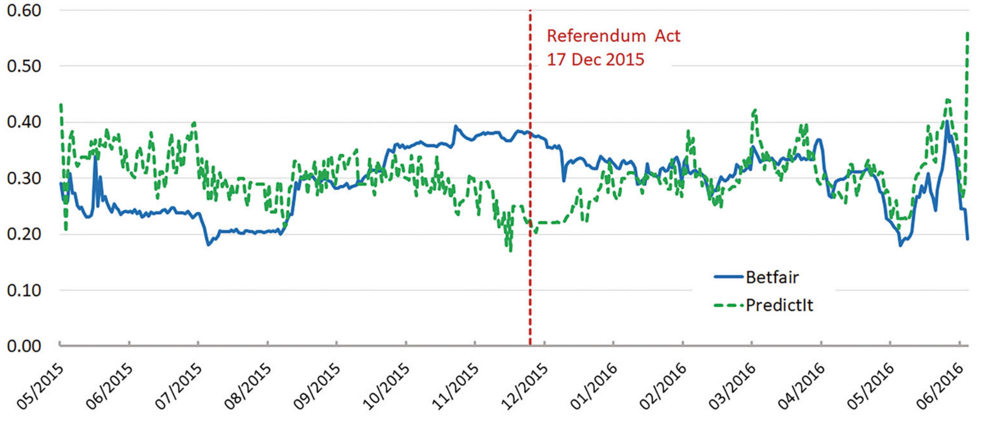

FIGURE 1 Odds of Leave. *Notes*: The figure shows the daily odds-derived probabilities of Leave provided by Betfair and PredictIt from 27 May 2015 to 23 June 2016. The dashed red vertical line identifies the date (17 December 2015) on which the Referendum Act received Royal Assent and thus came into force. Missing values are interpolated using previous-day values. The result of the vote became known in the early hours of 24 June 2016 in UK time, corresponding to the late hours of 23 June in US time; this explains the divergence of the two series on 23 June.

Among a series of relevant political events which marked the run up to the referendum and the progress of UK–EU negotiations, two dates were pivotal, according to the briefing paper by the UK House of Commons Library on Brexit (Walker 2018): the first is 17 December 2015—when the EU Referendum Act was promulgated—and the second is 22 February 2016—when Prime Minister David Cameron announced that a referendum would take place on 23 June 2016. The Referendum Act was the act of Parliament that made legal provision for a consultative referendum to be held in the UK (and Gibraltar), on whether it should remain a member state of the European Union or leave it. Following the Royal Assent to the Act, although the Prime Minister did not indicate a precise date for the vote, the British media considered June as the most likely period,[14] well before David Cameron's official announcement.[15] A dashed vertical line in Figure 1 identifies the date (17 December 2015) on which the European Union Referendum Act received Royal Assent and was therefore promulgated.

Figure 2 plots the average of the two Leave probability series (solid blue line, left-hand scale) along with the effective exchange rate of the BP against the basket of ten currencies (dashed yellow line, right-hand scale), using financial weights. Interestingly, around the approval of the EU Referendum Act, the exchange rate exhibits an upward movement (a depreciation of the BP of approximately 10%), while the correlation between the two variables increases—it is 0.18 from 27 May to 17 December 2015, and rises to 0.57 in the subsequent period. This suggests that from late 2015, the odds associated with the bets on the referendum result may have played a stronger role in explaining the BP exchange rate movements, consistently with the implication of our model. In subsection 5.3.6, we perform a number of additional tests. First, we weight the two measures with the respective daily number of bets (we do not observe the value of bets). We also use each measure separately, and we try a smoothed version of the daily volumes (see the discussion later). In all cases, the estimates remain robust.

Figure 3 plots all the (log) exchange rates considered in this paper, from 27 May 2015 to 23 June 2017. The financially weighted basket and the trade weighted basket are denoted by ROW^F and ROW^T, respectively (where 'ROW' stands for 'rest of the world'). The figure shows that, qualitatively, the BP exhibited similar patterns with respect to the different currencies considered in the paper, with large depreciations occurring around the Referendum Act of mid-December 2015 and after the Leave victory in the Referendum of June 2016.

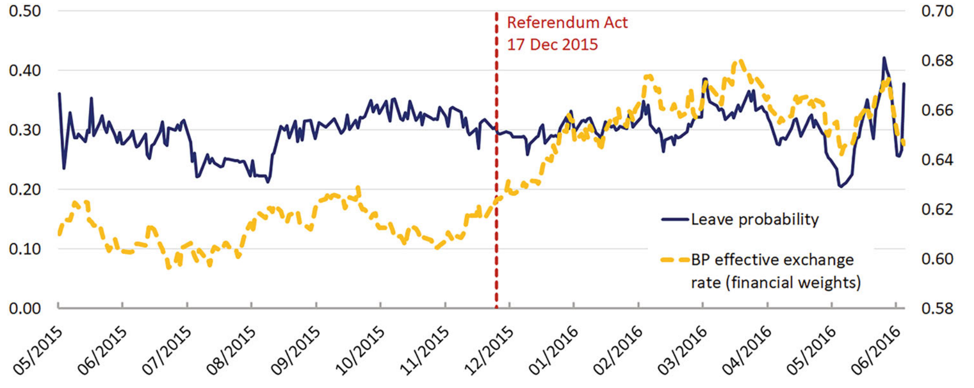

FIGURE 2 Brexit odds and the British pound. *Notes*: The figure shows our Leave probability measure (left-hand scale), constructed as the average of the Betfair and PredictIt odds-derived probabilities of Leave, along with the effective exchange rate of the BP vis-à-vis the basket of 10 currencies considered in the paper (right-hand scale). The basket is constructed as the weighted average of bilateral exchange rates using international financial positions as weights. The data are daily and span the period from 27 May 2015 to 23 June 2016. The dashed red vertical line identifies the date (17 December 2015) on which the Referendum Act received Royal Assent.

The additional measures of political and economic uncertainty that we collected will be introduced in the robustness section, as they are not employed in our baseline estimation.

5 | EMPIRICAL ANALYSIS

In this section, we estimate our model, which emphasizes the role played by market expectations and by political risk. For this purpose, we focus on the equilibrium (cointegration) relationship between exchange rates and their determinants, leaving aside short-run adjustment issues.[16]

5.1 | Setup

Based on equation (14), the effects of both the Leave probability and the political risk premium depend on market expectations.

We first estimate equation (14) for the effective (log) exchange rate of the BP vis-à-vis our two baskets of ten currencies (with either financial or trade weights). Next, we consider country-by-country equations for bilateral exchange rates. Finally, we estimate jointly a multi-currency portfolio model, which allows for cross-currency shock correlation and cross-equation constraints.

Our baseline sample spans the period from 27 May 2015, which is the first day when the odds on the referendum outcome are available, to 30 June 2016, that is, just a few days after the vote. However, we also estimate the equations on a longer sample ending in June 2017, to make sure that our main results do not simply depend on the specific end date of the baseline sample. Although we do not observe directly the *ex ante* probability q that a referendum will take place—there were no betting markets on this—we assume that this parameter is constant at q_0 before the Referendum Act, and takes value 1 after the date of promulgation of the Act. The value of q_0 will be estimated.

We provide the results of unit root and cointegration tests in the Online Appendix. The results broadly support our setup.

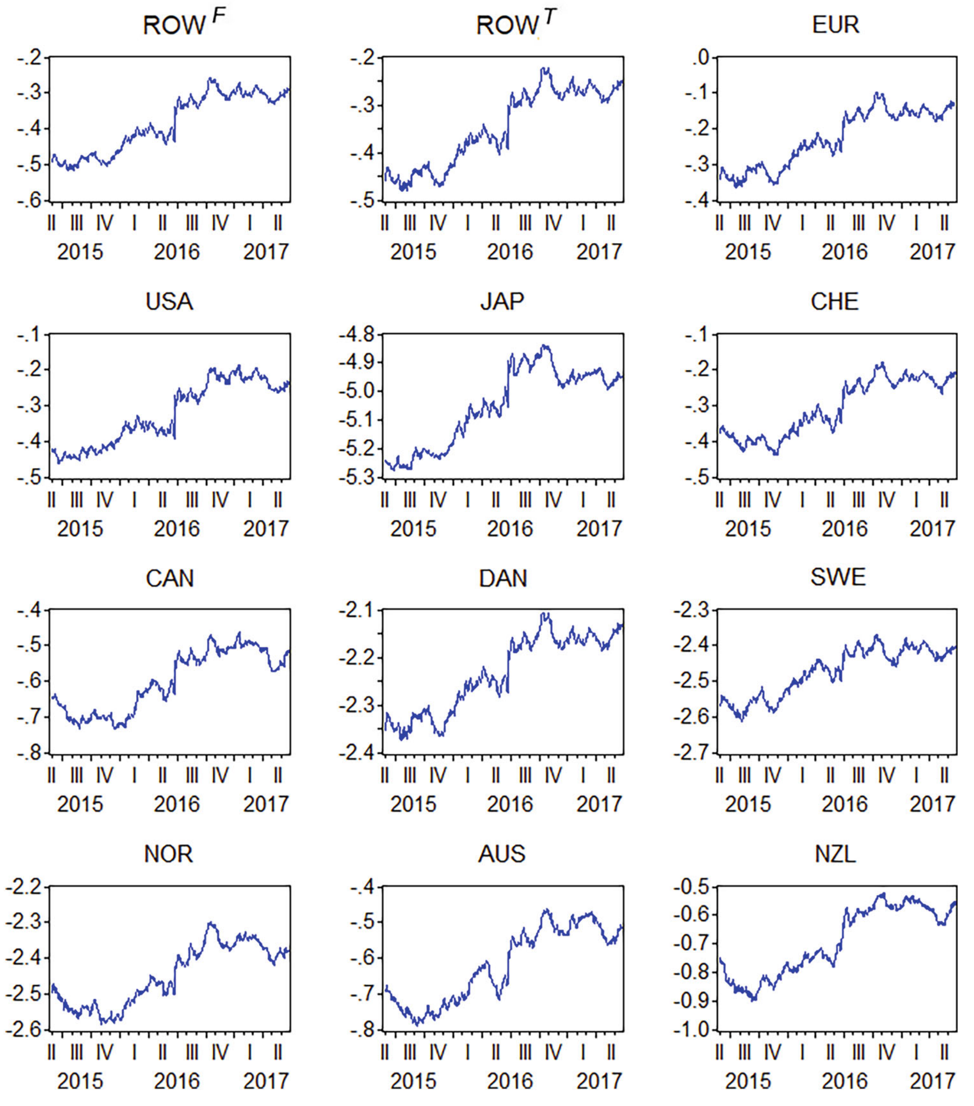

FIGURE 3 Log exchange rates of the British pound. *Notes*: The figure plots all log exchange rates used in the paper, from 27 May 2015 to 23 June 2017. An increase in any exchange rate means a depreciation of the BP against the currency considered. The exchange rates are displayed for the financially weighted basket (ROW^F) and the trade weighted basket (ROW^T), as well as all ten individual currencies: the euro (EUR), the US dollar (USA), the Japanese yen (JAP), the Swiss franc (CHE), the Canadian dollar (CAN), the Danish krone (DAN), the Swedish krone (SWE), the Norwegian krone (NOR), the Australian dollar (AUS) and the New Zealand dollar (NZL). The data are daily.

5.2 | Estimation results

This section presents our main results. We begin by considering ordinary least squares (OLS) estimates, with the (log) exchange rate as dependent variable. The OLS estimator is 'superconsistent' (i.e. it converges in probability to the true parameter value at speed T, the sample size, rather than the usual $\sqrt{T}$) when the variables are non-stationary but cointegrated. However, inference based on OLS standard errors is incorrect (Stock 1987). Moreover, Banerjee *et al.* (1986) show that the small sample bias of the OLS estimator can be substantial, and suggest estimating cointegration parameters through dynamic regressions rather than static regressions. In addition, Maddala and

Kim (1998) review the finite sample evidence on estimators of cointegrating vectors provided in Monte Carlo studies, and advise against estimating long-run parameters by static regressions.

For these reasons, we turn to the dynamic OLS (DOLS) estimator, proposed by Saikkonen (1991) and Stock and Watson (1993). We also show results obtained using the fully modified OLS (FMOLS) proposed by Phillips and Hansen (1990). Unlike the OLS, these estimators allow for inference on the coefficients of $I(1)$ variables in the presence of cointegration. As highlighted by Rossi (2013), cointegration vectors are typically estimated by DOLS in the literature on exchange rates.

The DOLS estimator applies a parametric correction to the OLS in order to account for the correlation between residuals and regressors. In practice, the estimator is obtained by augmenting the cointegration regression with lags and leads of first-differenced regressors. More specifically, in our case, the benchmark DOLS regression is given by

$$e_t = \theta(i_t^* - i_t) + \alpha + \beta q_t \pi_t + \gamma q_t \pi_t (1 - \pi_t) + \delta q_t + \sum_{j=-l}^{h} \boldsymbol{\phi}_j' \, \Delta \boldsymbol{X}_{t+j} + \varepsilon_t, \tag{16}$$

where $\boldsymbol{X}_{t+j} \equiv [i_{t+j}^* - i_{t+j}, q_{t+j}\pi_{t+j}, q_{t+j}\pi_{t+j}(1 - \pi_{t+j})]'$, $\Delta \boldsymbol{X}_{t+j} = \boldsymbol{X}_{t+j} - \boldsymbol{X}_{t+j-1}$, $\boldsymbol{\phi}_j$ is a 3×1 vector of parameters, for every j, and ε_t is an error term. An automatic lag/lead selection based on the Bayesian information criterion suggests including only contemporaneous differences in the regression, regardless of the value of q_0. However, we choose to also include lags and leads of order 1 (i.e. $l = h = 1$), as the results of F-tests indicate their overall significance.

As already mentioned, we also consider the FMOLS estimator (Phillips and Hansen 1990), which controls for the correlation between the error term of the cointegrating regression and the innovations of the regressors using a non-parametric consistent estimate of the long-run covariance matrix.

Table 2 reports the estimates obtained in the shorter sample 27 May 2015 to 30 June 2016 by DOLS, FMOLS and OLS for the financially weighted effective exchange rate (ROW^F) and the trade weighted effective exchange rate (ROW^T). To estimate q_0, we actually rely on non-linear least squares (however, to facilitate reading, we keep using the same labels as for ordinary least squares, namely, DOLS, FMOLS and OLS).[17]

Table 2 shows that the estimated coefficients of the Leave probability and the political risk premium, β and γ, respectively, are both positive and highly significant across all estimation methods. In our interpretation, β measures the percentage BP depreciation rate of the market prices in the Leave scenario, relative to the Remain scenario. The value of the BP conditional on Leave is estimated to be around 19%–22% lower than the value under the Remain scenario. Importantly, a positive coefficient for our measure of time-varying political risk premium, $\pi(1 - \pi)$, means that higher Brexit uncertainty is associated with a BP depreciation.

More generally, our approach allows us to disentangle the two channels (first and second moments) by which the odds affect the exchange rate. For instance, let us consider what the model predicts should happen to the BP after the Leave victory, using the estimates in column (1) of Table 2. Given the average values π (0.3) and $\pi(1 - \pi)$ (0.21) before the referendum, the model predicts a BP depreciation of about 7%. This is the net effect of: (i) the surprise of the referendum outcome, which accounts for a depreciation of $14.8\% = (1 - 0.3) \times 0.211$; (ii) an appreciation effect due to the resolution of uncertainty, $-8\% = (0 - 0.21) \times 0.384$. The net result is a depreciation rate of approximately 6.8%. Unlike a large body of the literature that finds negative coefficients on interest rate differentials, Table 2 shows that we obtain the *a priori* correct positive signs in our estimates for θ, although point estimates are generally higher than 1, and not significantly different from 0. However, we find that the regression for the basket of currencies does not violate the UIP, as the theoretical value $\theta = 1$ cannot be rejected. These results are consistent with those of a recent literature that also finds large positive UIP coefficients. For the

TABLE 2 DOLS, FMOLS and OLS estimates for the BP effective exchange rates, baseline sample (27/5/2015–30/6/2016).

	DOLS		FMOLS		OLS	
	ROW^F	ROW^T	ROW^F	ROW^T	ROW^F	ROW^T
	(1)	(2)	(3)	(4)	(5)	(6)
$q\pi$	0.211***	0.222***	0.208***	0.221***	0.185***	0.195***
	(0.033)	(0.038)	(0.035)	(0.041)	(0.030)	(0.035)
$q\pi(1-\pi)$	0.384***	0.420***	0.369***	0.417***	0.282***	0.320***
	(0.101)	(0.119)	(0.114)	(0.134)	(0.091)	(0.107)
$i^* - i$	2.222	2.913	1.757	2.659	0.613	1.150
	(2.709)	(2.504)	(2.300)	(2.408)	(1.409)	(1.604)
Constant	−0.502***	−0.450***	−0.501***	−0.444***	−0.507***	−0.455***
	(0.023)	(0.023)	(0.012)	(0.016)	(0.024)	(0.022)
q	−0.052	−0.065	−0.052	−0.073*	−0.027	−0.045
	(0.039)	(0.043)	(0.034)	(0.039)	(0.042)	(0.043)
q_0	0.254	0.198	0.227	0.119	0.228	0.135
	(0.170)	(0.196)	(0.223)	(0.236)	(0.198)	(0.218)

Notes: The table reports non-linear DOLS, FMOLS and OLS estimates (and standard errors in parentheses) of the equations for the (log) effective exchange rates of the BP against the baskets of ten currencies considered in the paper (ROW^F for financially weighted basket, ROW^T for trade-weighted basket). See note 17 for more details on the non-linear methods employed. Here, π is the Leave probability, q is the probability of the referendum being held, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates, q takes value q_0 before the Referendum Act and 1 afterwards. The corresponding line in the table reports the estimated coefficients on this dummy variable, while the bottom part of the table reports the estimates of q_0. The reported standard errors for DOLS and OLS are heteroscedasticity- and autocorrelation-consistent (HAC). In FMOLS, the long-run error variance is estimated by the Bartlett kernel with Newey–West fixed bandwidth of 6. The FMOLS standard error of q_0 is calculated as $\hat{\sigma}_{lr}(\tilde{h}/2)^{-1/2}$, where $\hat{\sigma}_{lr}$ is the long-run standard error of FMOLS residuals, and $\tilde{h}$ is the second derivative of the sum of squared residuals with respect to q_0. For the other parameters, the FMOLS standard errors are conventional FMOLS standard errors, conditional on q_0.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

BP, Bussiere *et al.* (2022, Table 1, panel C) find a UIP coefficient around 5.3 over the sample 2005M05–2017M04, while Engel *et al.* (2022, Table 4) obtain an estimate around 4.6 over the sample 2007M01–2020M09, with wide confidence intervals, similarly to our results. However, as shown below, when we extend our sample, we obtain estimates that are much closer to the theoretical value 1.

The DOLS estimates of q_0 are 0.25 and 0.2 for ROW^F and ROW^T, respectively, with large confidence intervals (at 90%, the parameter lies between 0 and 0.6, approximately). The estimates are somewhat lower for FMOLS and OLS.[18] Thus, as expected, the effects of the odds variables are much higher after the Referendum Act, which confirms the intuition that only when market participants perceive that the referendum will take place for sure do they start placing more weight on the evolution of the odds.[19] Also, the coefficient δ on the variable q (non-interacted) is not significantly different from zero, except for FMOLS estimates for ROW^T. This is consistent with the assumption that market participants expected the same value for the exchange rate in the case of No Referendum and in the case of a Remain victory. If the non-interacted q variable is omitted from the regression, then the estimates of q_0 become larger and significant, but still remain below 0.5 (the DOLS point estimates are around 0.44 for ROW^F, and 0.43 for ROW^T).

In our interpretation, the term $\alpha + \delta q$ gives the expected (log) exchange rate conditional on either a Remain vote or no referendum at all. This implies that after the Referendum Act ($q = 1$), the expected multilateral (financially weighted) BP exchange rate conditional on a Remain victory is approximately equal to $\exp(\alpha + \delta) = \exp(-0.502 - 0.052) = 0.57$, which is close to but below the average rate prevailing before the Referendum Act (see Figure 2).

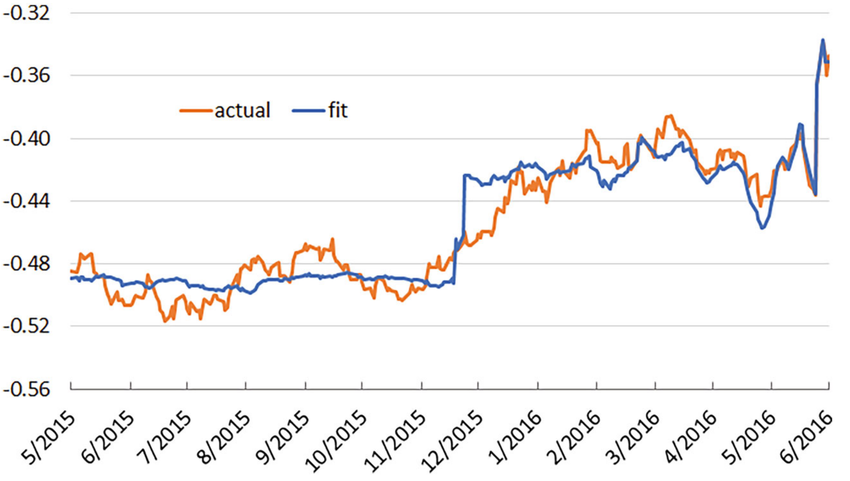

FIGURE 4 Fit of the model. *Notes*: The figure compares the fitted values for ROW^F based on the DOLS estimates in Table 2 with the actual series of the ROW^F log exchange rate. The sample goes from 27 May 2015 to 30 June 2016.

TABLE 3 DOLS estimates for bilateral exchange rates of the British pound, baseline sample (27/5/2015–30/6/2016).

	EUR	USA	JAP	CHE	CAN
$q\pi$	0.227***	0.172***	0.301***	0.259***	0.217***
	(0.043)	(0.018)	(0.081)	(0.034)	(0.054)
$q\pi(1-\pi)$	0.441***	0.316***	0.217	0.462***	0.465***
	(0.133)	(0.057)	(0.255)	(0.105)	(0.172)
$i^* - i$	2.863	5.122***	−12.699***	3.184	17.447***
	(2.334)	(1.700)	(3.373)	(2.973)	(1.921)
Constant	−0.332***	−0.437***	−5.373***	−0.367***	−0.716***
	(0.018)	(0.006)	(0.024)	(0.045)	(0.003)
q	−0.057	−0.047**	0.023	−0.111***	−0.097*
	(0.044)	(0.020)	(0.081)	(0.034)	(0.055)

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses) for the bilateral log exchange rates of the BP against the euro (EUR), the US dollar (USA), the Japanese yen (JAP), the Swiss franc (CHE) and the Canadian dollar (CAN). Here, π is the Leave probability, q is the probability of the referendum being held, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates; q takes value $q_0 = 0.25$ before the Referendum Act and 1 afterwards.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

Figure 4 shows the model fit for the financially weighted multilateral exchange rate, based on the DOLS estimates from Table 2.

Tables 3 and 4 report the DOLS estimates for each individual currency. Here, we impose the *a priori* restriction that q_0 is the same for all countries, and equal to 0.25.

The coefficient on the Leave probability $q\pi$ is positive and significant at the 1% level for all currencies, with point estimates ranging from 0.17 to 0.30. The coefficient on $q\pi(1-\pi)$ is also positive for all countries, and significant at 1% for all countries except Japan (for which it is not significant), Norway (for which it is significant at 10%) and New Zealand (for which it is significant at 5%). The estimates for the interest rate spread coefficient θ exhibit large cross-sectional variability. However, for all countries except Japan and Sweden, the point estimate is positive. Due to the large standard errors, the coefficient is not significantly different from 0 (or 1) in the equations for the euro, the Swiss franc, the Swedish krone, the Australian dollar and the New Zealand dollar. For the other currencies, the UIP condition is more clearly violated.

TABLE 4 More DOLS estimates for bilateral exchange rates of the British pound, baseline sample (27/5/2015–30/6/2016).

	DAN	SWE	NOR	AUS	NZL
$q\pi$	0.230***	0.225***	0.216***	0.304***	0.302***
	(0.047)	(0.046)	(0.067)	(0.052)	(0.050)
$q\pi(1-\pi)$	0.441***	0.497***	0.406*	0.594***	0.428**
	(0.146)	(0.148)	(0.207)	(0.167)	(0.167)
$i^* - i$	5.122***	−0.659	9.142**	0.107	2.629
	(1.416)	(2.190)	(3.593)	(3.741)	(2.813)
Constant	−2.305***	−2.599***	−2.604***	−0.770***	−0.941***
	(0.017)	(0.024)	(0.019)	(0.066)	(0.075)
q	−0.067	−0.068	−0.073	−0.126**	−0.050
	(0.048)	(0.046)	(0.068)	(0.053)	(0.048)

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses) for the bilateral log exchange rates of the BP against the Danish krone (DAN), the Swedish krone (SWE), the Norwegian krone (NOR), the Australian dollar (AUS) and the New Zealand dollar (NZL). Here, π is the Leave probability, q is the probability of the referendum being held, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates; q takes value $q_0 = 0.25$ before the Referendum Act and 1 afterwards.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

TABLE 5 DOLS, FMOLS and OLS estimates for the BP effective exchange rates, long sample (27/5/2015–23/6/2017).

	DOLS		FMOLS		OLS	
	ROW^F	ROW^T	ROW^F	ROW^T	ROW^F	ROW^T
$q\pi$	0.249***	0.254***	0.243***	0.252***	0.220***	0.227***
	(0.033)	(0.038)	(0.041)	(0.043)	(0.030)	(0.035)
$q\pi(1-\pi)$	0.295***	0.342***	0.273**	0.333**	0.197**	0.243**
	(0.104)	(0.120)	(0.136)	(0.142)	(0.092)	(0.106)
$i^* - i$	1.189	2.153	1.030	2.044	0.967	1.467
	(2.23)	(2.248)	(1.770)	(1.950)	(1.650)	(1.657)
Constant	−0.505***	−0.452***	−0.507***	−0.451***	−0.506***	−0.454***
	(0.021)	(0.022)	(0.010)	(0.013)	(0.024)	(0.022)
q	−0.047	−0.062	−0.041	−0.063	−0.019	−0.038
	(0.038)	(0.042)	(0.042)	(0.043)	(0.042)	(0.043)
q_0	0.231	0.171	0.245	0.147	0.232	0.141
	(0.170)	(0.198)	(0.286)	(0.265)	(0.197)	(0.217)

Notes: See Table 2.

As shown in Tables 5–7, our main results remain valid when we extend the sample by including one year of post-referendum observations. In all equations, the estimates of the Leave probability coefficient, β, are larger, and those of the risk premium coefficient, γ, are lower, than over the short sample, but in general the estimates are very similar. Moreover, the estimates of the interest spread coefficient, θ, are now much closer to the theoretical value 1 both for the basket of currencies and for individual currencies, with the exception of the Swiss franc and the Australian dollar. Notably, the interest spread coefficient is now not significantly different from 1 for the US dollar either.

TABLE 6 DOLS estimates for bilateral exchange rates of the British pound, long sample (27/5/2015–23/6/2017).

	EUR	USA	JAP	CHE	CAN
$q\pi$	0.256***	0.228***	0.343***	0.281***	0.231***
	(0.042)	(0.022)	(0.085)	(0.031)	(0.059)
$q\pi(1-\pi)$	0.376***	0.197***	0.244	0.417***	0.428**
	(0.133)	(0.067)	(0.274)	(0.104)	(0.187)
$i^* - i$	2.763	1.893	−9.189***	4.240**	12.655***
	(2.005)	(1.705)	(2.293)	(1.812)	(2.005)
Constant	−0.333***	−0.447***	−5.349***	−0.350***	−0.714***
	(0.016)	(0.006)	(0.017)	(0.028)	(0.003)
q	−0.053	−0.029	0.015	−0.109***	−0.096
	(0.044)	(0.022)	(0.087)	(0.033)	(0.060)

Notes: See Table 3.

TABLE 7 More DOLS estimates for bilateral exchange rates of the British pound, long sample (27/5/2015–23/6/2017).

	DAN	SWE	NOR	AUS	NZL
$q\pi$	0.254***	0.243***	0.248***	0.369***	0.343***
	(0.046)	(0.047)	(0.066)	(0.048)	(0.045)
$q\pi(1-\pi)$	0.386***	0.478***	0.323	0.475***	0.283*
	(0.145)	(0.151)	(0.206)	(0.161)	(0.149)
$i^* - i$	4.962***	0.904	8.082***	−1.937	0.826
	(1.301)	(1.689)	(2.512)	(2.419)	(2.108)
Constant	−2.307***	−2.584***	−2.598***	−0.733***	−0.892***
	(0.015)	(0.018)	(0.014)	(0.043)	(0.055)
q	−0.063	−0.065	−0.068	−0.124**	−0.045
	(0.047)	(0.047)	(0.068)	(0.049)	(0.044)

Notes: See Table 4.

5.3 | Robustness checks for single-equation estimation

5.3.1 | Unrestricted break model

In our benchmark equation, we have posited that a structural break occurs around the promulgation of the EU Referendum Act. This subsection shows that this assumption is supported by the data and is largely inconsequential for the estimation results.

We estimate the basic model (11) by DOLS and perform the Quandt–Andrews test for parameter instability at one unknown break date. We thus allow the data to determine if and when the coefficients of π and $\pi(1-\pi)$, that is, β and γ, exhibit structural breaks. We use a 'pre'/'post' suffix to denote the parameter estimates for the pre/post break period.

Table 8 shows the results for the multilateral exchange rates ROW^F and ROW^T. For both the trade and financially weighted rates, we strongly reject the hypothesis that the coefficients do not change during the sample period. In both equations, we estimate a break occurring on 7 January 2016—only a few weeks after our prior of a break, 17 December 2015. Importantly, in the post-break sample, the estimated coefficients of the Leave probability and the risk premium, β_{post} and γ_{post}, are strongly significant and not statistically different from our benchmark estimates for both exchange rates (see Table 2), while the coefficients are not significantly different from

TABLE 8 Robustness: unrestricted breaks in β and γ.

Parameter	ROW^F	ROW^T
β_{pre}	−0.061	−0.151
	(0.214)	(0.260)
γ_{pre}	0.337	0.484
	(0.382)	(0.456)
β_{post}	0.183***	0.192***
	(0.023)	(0.027)
γ_{post}	0.325***	0.345***
	(0.078)	(0.091)
θ	3.815*	2.911
	(2.273)	(2.303)
α	−0.520***	−0.487***
	(0.023)	(0.030)
Break date	7/1/2016	7/1/2016
F-test $\beta_{pre} = \gamma_{pre} = 0$	0.0496	0.1178

Notes: The table shows the DOLS estimates (and HAC standard errors in parentheses) of equation (11), with breaks in coefficients β and γ at an unknown date. Results are reported for the effective exchange rates ROW^F and ROW^T. Here, β_{pre} and β_{post} denote the values of β (i.e. the coefficient of variable π) before and after the promulgation of the Referendum Act, respectively; γ_{pre} and γ_{post} denote the values of γ (i.e. the coefficient of variable $\pi(1-\pi)$) before and after the promulgation of the Referendum Act, respectively; θ is the coefficient of the interest rate differential $i^* - i$; and α is the constant. The bottom part of the table reports the estimated break date and the p-value of an F-test of joint non-significance of β_{pre} and γ_{pre}. The sample goes from 27 May 2015 to 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

zero in the pre-break period. Note, however, that this can be due to the fact that the covariates π and $\pi(1-\pi)$ in the pre-break subsample are highly collinear, and this may impair inference about their individual coefficients.

An F-test with the null hypothesis that the Leave probability and the risk premium are jointly non-significant before the break has a p-value slightly below 5% in the case of ROW^F, and about 12% in the case of ROW^T.

Overall, these results confirm that our assumption of a structural break occurring at the time of the Referendum Act is consistent with the data, and that the assumption does not drive the estimates of the other coefficients.

5.3.2 | Alternative uncertainty measures

This subsection compares our model-based measure of political risk premium with four standard measures of financial and economic uncertainty. In particular, we show here that the estimated effect of our measure based on the bookmakers' Leave probability is robust to their inclusion.

The first measure that we consider is the Economic Policy Uncertainty (EPU) index for the UK developed by Baker *et al.* (2016). The index is constructed by counting newspaper articles that contain at least one term from each of three sets of keywords: (i) economic or economy; (ii) uncertain or uncertainty; (iii) the policy-related words spending, deficit, regulation, budget, tax, policy, and Bank of England. The index covers the 650 UK newspapers included in the digital archives of the Access World News NewsBank service. The data are provided at http://www.policyuncertainty.com.

The other three measures capture financial uncertainty and are derived from option markets. The second measure is the UK counterpart of the VIX index, or VFTSE, which is

TABLE 9 Correlations between uncertainty measures.

	$\pi(1-\pi)$	epu_{uk}	$vftse$	vol_{fx}	$riskrev$
$\pi(1-\pi)$	1				
	—				
epu_{uk}	0.150**	1			
	(2.539)	—			
$vftse$	0.021	0.211***	1		
	(0.359)	(3.620)	—		
vol_{fx}	0.259***	0.564***	0.101*	1	
	(4.486)	(11.434)	(1.696)	—	
$riskrev$	0.167***	0.540***	−0.004	0.964***	1
	(2.827)	(10.730)	(−0.064)	(61.327)	—

Notes: The table reports the correlation coefficients between the different uncertainty measures considered in our robustness checks, along with the associated t-statistics (in parentheses). Here, $\pi(1-\pi)$ is our measure of odds-implied political risk; epu_{uk} is the Economic Policy Uncertainty index (Baker *et al.* 2016) for the UK, divided by 100; $vftse$ is the VFTSE index, i.e. the option-implied 1-month-ahead volatility of the FTSE 100 stock market index; vol_{fx} is the index of option-implied 3-month-ahead volatility of the BP/USD exchange rate; and $riskrev$ is the risk reversal of the BP, i.e. the difference between the implied volatility of out-of-the-money put options and the implied volatility of symmetric out-of-the-money call options. All correlations are computed over the sample 27 May 2015 to 23 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

the option-implied 1-month-ahead volatility in the FTSE 100 equity index. It represents the (risk-neutral) expected standard deviation of the stock market index. The third measure is the risk-neutral expected volatility of the BP/USD exchange rate implied by the prices of exchange rate options with 3-month horizons. The fourth measure is the so-called 'risk reversal' of the BP. A risk reversal is the difference between the implied volatility of out-of-the-money put options and the implied volatility of symmetric out-of-the-money call options.[20] This difference measures the premium paid for protection against the expected (skewness towards) depreciation in the distribution of the foreign currency. A recent literature that focuses on disaster risk as a key determinant of the exchange rate (e.g. Brunnermeier *et al.* 2008; Farhi and Gabaix 2015) has underlined a close relationship between the risk reversal and the level of the exchange rate. All three option market variables actually capture a combination of uncertainty and risk aversion; Bekaert *et al.* (2013) provide estimates of the two components in the case of the VIX. The data source for all three series is Datastream. Table 9 shows that the correlations between different uncertainty measures and our odds-implied volatility are positive, statistically significant, and of magnitude between 0.15 and 0.26, except for the VFTSE.

Table 10 reports the results of our robustness regression for the financially weighted multilateral exchange rate, using DOLS. Columns (1)–(4) show the results obtained when the additional regressors are included one at a time. Because the exchange rate option-implied volatility and the risk reversal are extremely highly correlated (their correlation coefficient is 0.96), a regression with all uncertainty measures together suffers from severe collinearity problems. Thus in column (5), we include all uncertainty measures except the risk reversal. The table shows that the coefficients of our measures of the Leave probability and of the risk premium, $q\pi$ and $q\pi(1-\pi)$, respectively, remain highly significant irrespective of the inclusion of the other volatility indicators. All the uncertainty indicators, with the exception of the VFTSE, when included one by one are significant and have the expected positive sign, implying that more uncertainty is associated with a weaker BP. Interestingly, when the risk reversal measure is included in the equation, the point estimate of the UIP coefficient is very close to the theoretical value 1.

TABLE 10 Robustness: other uncertainty measures.

	(1)	(2)	(3)	(4)	(5)
$q\pi$	0.1421***	0.2048***	0.1593***	0.1946***	0.1275***
	(0.0350)	(0.0341)	(0.0173)	(0.0207)	(0.0254)
$q\pi(1-\pi)$	0.4740***	0.3806***	0.2299***	0.2977***	0.1846***
	(0.1082)	(0.1082)	(0.0681)	(0.0703)	(0.0873)
$i^* - i$	2.9494	2.0890	0.6785	1.0846	0.2208
	(2.7274)	(2.5178)	(2.7618)	(2.7956)	(2.5717)
Constant	−0.5051***	−0.5108***	−0.5315***	−0.5074***	−0.5540***
	(0.0152)	(0.0159)	(0.0184)	(0.0155)	(0.0214)
q	−0.0574*	−0.0512	−0.0248	−0.0413*	−0.0132
	(0.0302)	(0.0351)	(0.0201)	(0.0232)	(0.0202)
epu_{uk}	0.0045**				0.0006
	(0.0019)				(0.0019)
$vftse$		0.0005			0.0009**
		(0.0004)			(0.0004)
vol_{fx}			0.0035***		0.0040***
			(0.0009)		(0.0011)
$riskrev$				0.0039***	
				(0.0015)	

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses). The dependent variable is the log financially weighted effective exchange rate of the BP against the basket of ten currencies considered in the paper. Here, π is the Leave probability, q is the probability of the referendum being held, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates; q takes value $q_0 = 0.25$ before the Referendum Act and 1 afterwards. Also, epu_{uk} is the Economic Policy Uncertainty index (Baker *et al.* 2016) for the UK, divided by 100; $vftse$ is the VFTSE index, i.e. the option-implied 1-month-ahead volatility of the FTSE 100 stock market index; vol_{fx} is the index of option-implied 3-month-ahead volatility of the BP/USD exchange rate; and $riskrev$ is the risk reversal of the BP, i.e. the difference between the implied volatility of out-of-the-money put options and the implied volatility of symmetric out-of-the-money call options. The sample goes from 27 May 2015 to 30 June 2016. In column (5), $riskrev$ is excluded to avoid collinearity problems (the correlation with vol_{fx} is 0.96).
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

5.3.3 | Allowing for time-varying risk aversion

Our baseline results have been obtained under the assumption of a time-invariant risk aversion parameter in investors' utility function. In this subsection, we show that our main findings on the relationship between BP exchange rates and political risk are robust to the inclusion of two different measures of time-varying risk aversion.

The first measure (which we label as ra_t) is derived from data on option-implied and realized volatilities, following the popular approach proposed by Bekaert *et al.* (2013). These authors decompose option-implied volatility in the US stock market (the VIX index) into an uncertainty component and a risk aversion component. We apply their methodology to measure risk aversion using data on foreign-exchange market volatility.[21] To be consistent with the 3-month maturity considered for interest rates in our baseline regressions, we construct a time series of risk aversion by decomposing 3-month (65-trading-day) option-implied volatility (in their paper, Bekaert *et al.* 2013 decompose 1-month, 22-trading-day, volatility). However, our main results are also robust to the use of 1-month volatility. We proceed as follows.

First, we use hourly data on exchange rates and calculate the hour-on-hour percentage changes in a currency exchange rate vis-à-vis the BP (the hourly return). The (cumulated) realized variance of the BP exchange rate against each currency at day t, $rvar_t$, is calculated by summing the squared hourly changes over the previous 65 trading days (including day t itself). We regress

TABLE 11 Robustness: time-varying risk aversion.

	ROW^F	ROW^F	ROW^T	ROW^T
$q\pi$	0.213***	0.203***	0.225***	0.214***
	(0.029)	(0.025)	(0.033)	(0.029)
$q\pi(1-\pi)$	0.287***	0.380***	0.295***	0.413***
	(0.102)	(0.083)	(0.112)	(0.094)
$i^* - i$	1.935	2.825	3.465	3.153
	(2.664)	(2.506)	(2.441)	(2.249)
Constant	−0.504***	−0.513***	−0.454***	−0.466***
	(0.015)	(0.024)	(0.016)	(0.030)
q	−0.039	−0.045*	−0.043	−0.051*
	(0.030)	(0.026)	(0.034)	(0.029)
ra	−1.243		−1.780*	
	(0.814)		(0.910)	
bex		0.440		0.322
		(0.619)		(0.813)

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses). The dependent variable is either the log financially weighted (ROW^F) or the log trade weighted (ROW^T) effective exchange rate of the BP against the basket of ten currencies considered in the paper. Here, π is the Leave probability, q is the probability of the referendum being held, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates; q takes value $q_0 = 0.25$ before the Referendum Act and 1 afterwards. Also, ra denotes time-varying risk aversion estimated from realized and option-implied volatilities in the foreign exchange market; and bex is the index of time-varying risk aversion by Bekaert *et al.* (2022) (divided by 100). The sample goes from 27 May 2015 to 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

this variable on the time $t-65$ option-implied volatility for time t ($ivar_{t-65,t}$) and on the past realized variance, $rvar_{t-65}$, that is,

$$rvar_t = c_0 + c_1 \cdot ivar_{t-65,t} + c_2 \cdot rvar_{t-65} + \epsilon_t, \tag{17}$$

where c_0, c_1 and c_2 are coefficients, and ϵ_t is an error term. The proxy for the time-varying risk aversion associated with the currency/BP rate is measured as the excess of the implied volatility over the predicted value for realized variance from regression (17), that is,

$$ra_{t-65} = ivar_{t-65,t} - \widehat{rvar}_t. \tag{18}$$

To have an aggregate measure of risk aversion for the BP, we calculate a weighted average of *rvar* and *ivar* across different currencies, using the financial weights reported in Table 1, and estimate equation (17) on these averages.[22] We estimate ra_t over our baseline sample.

The second measure (bex_t) that we employ is the index of time-varying risk aversion developed recently by Bekaert *et al.* (2022) using a dynamic no-arbitrage asset pricing model. This we take 'off the shelf' for our sample period. This measure reflects the utility function of a representative agent in a generalized habit-like model with preference shocks, and is estimated through an instrumental variable approach using a set of observed financial variables (earnings yield, credit and term spreads, equity and corporate bond realized variances, and equity risk-neutral variance).

Table 11 reports the results of our regressions for multilateral exchange rates of the BP, augmented with the two measures of risk aversion. As the table shows, the estimated coefficients on referendum odds and odds-implied political risk remain positive and strongly significant.

TABLE 12 Other robustness checks.

Panel A:			Panel B:	
No referendum uncertainty			Weights based on currency exposures	
	ROW^F	ROW^F		ROW^C
π	0.209***	0.162***	$q\pi$	0.201***
	(0.041)	(0.011)		(0.029)
$\pi(1-\pi)$	0.330***	0.238***	$q\pi(1-\pi)$	0.373***
	(0.133)	(0.044)		(0.087)
i^*-i	−2.941	−1.246	i^*-i	3.709
	(4.605)	(2.673)		(2.713)
Constant	−0.571***	−0.519***	Constant	−0.439***
	(0.045)	(0.015)		(0.026)
			q	−0.041
				(0.037)
Sample start	18/12/2015	23/2/2016	q_0	0.315**
				(0.158)

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses). In panel A, the dependent variable is the log financially weighted (ROW^F) effective exchange rate of the BP against the basket of ten currencies considered in the paper. Here, π is the Leave probability, and i^*-i is the difference between the foreign and domestic (UK) interest rates. Two alternative sample start dates are considered: 18 December 2015, i.e. the day after the promulgation of the Referendum Act, and 23 February 2016, i.e. the day after the official announcement of the referendum date. The sample end date is 30 June 2016. In panel B, the dependent variable (labelled as ROW^C) is the log multilateral exchange rate of the BP against the US dollar, the euro and the Japanese yen, calculated as a weighted average using cross-border currency exposures (Benetrix *et al.* 2019) as weights. Here, π is the Leave probability; q is the probability of the referendum being held, taking value q_0 before the Referendum Act and 1 afterwards; i^*-i is the difference between the foreign (weighted by currency exposures) and domestic (UK) interest rates. The bottom part of the table reports the estimate of q_0. The sample goes from 27 May 2015 to 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

5.3.4 | Estimates with no referendum uncertainty

Next, we report the estimates obtained when the uncertainty on the occurrence of the referendum is removed. In particular, we restrict our estimation sample to: (i) the period following the promulgation of the Referendum Act (18 December 2015 to 30 June 2016); (ii) the period following the official announcement of the referendum date (23 February 2016 to 30 June 2016).

Panel A of Table 12 reports the results. The estimates of the coefficients on π and $\pi(1-\pi)$ remain strongly significant and very similar to those obtained over our baseline sample (see Table 2).

5.3.5 | Alternative currency weights: cross-border currency exposures

In this subsection, we consider an alternative set of weights to construct multilateral exchange rates and interest rates. Specifically, we use weights based on the currency composition of the international investment position of the UK, as provided by Benetrix *et al.* (2019) in their dataset on cross-border currency exposures of individual countries over the period 1990–2017. In particular, we consider the shares of the UK's external assets denominated in US dollars, euros and Japanese yen as of 2015, as well as the shares of the UK's external liabilities denominated in these currencies in the same year, and calculate currency weights by taking an average of asset shares and liability shares for each currency (these weights are then rescaled so that they sum to 1).[23]

Panel B of Table 12 reports the estimates of equation (14) based on these alternative measures of multilateral exchange rates and interest rates. The results are very similar to our baseline estimates from Table 2.

5.3.6 | Weighted averages of PredictIt and Betfair odds

In the analysis so far, our measure for the probability of Leave was calculated as the average of the PredictIt and Betfair probabilities. As a final robustness check, we consider different measures constructed using alternative weighting schemes between the two companies. The first set of weights is based on the number of daily bets (we do not observe their values) placed on each betting platform over the total number of bets observed on the two platforms. Of course, these weights change on a daily basis. On average, this scheme assigns PredictIt a weight of about 62%. The resulting weighted average turns out to be more volatile than the benchmark (equally weighted) average, for two reasons. The first is that it gives greater weight to the more volatile series, PredictIt. The second is that it introduces an additional source of variability, that of the weights. Since the daily number of bets is quite volatile between the two agencies, this possibly introduces some noise in the weighted measure of the Leave probability. Second, we use each individual series separately. We re-estimate the baseline model using these new definitions of Leave probability π (weighted average, only PredictIt, only Betfair). The results, for the baseline sample as well as the shortest sample with no referendum uncertainty introduced in subsection 5.3.4, are reported in Table 13. As discussed in Section IV, the odds provided by the two betting platforms exhibit some noise throughout 2015, when considered individually, whereas they are strongly aligned with each other from January 2016 onwards. Using the weighted average produces significant estimates for the coefficients on the Leave probability and the political risk variables in both samples (see columns (1) and (4)). The point estimates, compared to the benchmark regression of Table 2, are somewhat smaller, a feature that we attribute to the extra volatility introduced by variable weights. In fact, if we re-estimate the model with unequal but constant weights (62% to PredictIt and 38% to Betfair), or with weights based on a moving average of the daily volumes, then we actually obtain point estimates (not shown in the table) that are higher than in Table 2 and strongly significant. Looking at the results for the individual time series of PredictIt and Betfair, the two measures of odds provide similar results, especially in the shortest sample, that is, the one starting after the announcement of the referendum date on 22 February 2016. In particular, over this sample, the coefficients on odds and odds-implied volatility are strongly significant for both odds measures. Over the baseline sample, the coefficient on π is always significant, while the coefficient on the odds-implied volatility $\pi(1-\pi)$ is statistically significant when using PredictIt data, but it is not when using the Betfair probability.

5.4 | System estimation

Our portfolio approach suggests that efficiency gains can be obtained by exploiting the information contained in the co-movement of exchange rates. In fact, the model implies that as long as investors choose the asset/currency composition of their portfolios by considering interest yields and the covariance structure of currencies, their risk attitude should be the same across all currencies. This implies that all bilateral BP exchange rates should be affected equally by a change in the political risk premium associated with Brexit. In this subsection, we estimate a system of equations (16) simultaneously, and test the cross-equation restriction that the coefficient of the volatility term $q\pi(1-\pi)$ (i.e. the parameter γ) should be the same for all currencies. As happens with SUR and OLS estimators, the DOLS system and single-equation estimators coincide when cross-currency error terms are not correlated, while the system estimator, the dynamic

TABLE 13 Robustness: different weights on PredictIt and Betfair Leave probabilities.

	Weighted average (1)	PredictIt (2)	Betfair (3)	Weighted average (4)	PredictIt (5)	Betfair (6)
$q\pi$	0.152***	0.194***	0.105***	0.149***	0.141***	0.152***
	(0.028)	(0.021)	(0.030)	(0.014)	(0.016)	(0.01)
$q\pi(1-\pi)$	0.183*	0.326***	0.040	0.185***	0.153**	0.213***
	(0.101)	(0.067)	(0.125)	(0.052)	(0.059)	(0.039)
$i - i^*$	0.513	0.458	0.920	−2.088	−2.693	−1.614
	(2.978)	(2.690)	(3.494)	2.809	(2.974)	(2.222)
Constant	−0.511***	−0.512***	−0.509***	−0.509***	−0.504***	−0.510***
	(0.016)	(0.015)	(0.019)	(0.018)	(0.0205	(0.013)
q	0.006	−0.032	0.052*			
	(0.028)	(0.021)	(0.031)			
Sample start	27/5/2015			23/2/2016		

Notes: The table reports DOLS estimates (and HAC standard errors in parentheses). The dependent variable is the log financially weighted (ROW^F) effective exchange rate of the BP against the basket of ten currencies considered in the paper. Here, π denotes the Leave probability and is measured in three alternative ways: (i) as an average of the probabilities provided by PredictIt and Betfair, weighted by the daily number of bets placed in the two platforms; (ii) using the Leave probability provided by PredictIt only; (iii) using the Leave probability provided by Betfair only. Also, q denotes the probability of the referendum being held, taking value $q_0 = 0.25$ before the Referendum Act and 1 afterwards; $i^* - i$ is the difference between the foreign and domestic (UK) interest rates. The table reports results obtained over both our baseline sample, starting on 27 May 2015 (when odds data become available), and the shorter sample considered in panel A of Table 12, starting on 23 February 2016 (i.e. the day after the official announcement of the referendum date). The sample end date is 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

SUR (DSUR) (see Mark *et al.* 2005), is more efficient than the single-equation DOLS when the cross-currency errors are correlated. Like the static SUR, the DSUR is a two-step estimator: in the first step, the system is estimated by dynamic OLS in order to obtain a consistent estimate of the covariance matrix of the errors; in the second step, all the other model parameters are estimated by (dynamic) generalized least squares (GLS), taking the covariance matrix from step 1 as given. To account for autocorrelation, the DSUR approach considers the long-run covariance matrix of the error terms. As we consider a system of equations, we want to make sure that investors' prior expectations concerning the likelihood of a vote taking place, q_0, is the same irrespective of the bilateral exchange rate under consideration.

Imposing the cross-equation equality restriction on q_0 introduces another non-linear constraint into the system. To address this issue, we proceed as follows. In the first step of our DSUR estimation procedure, we run a nonlinear DOLS and estimate q_0, as well as the long-run covariance matrix of residuals,[24] imposing that q_0 is the same in all equations. In the second step, we use the estimate for q_0 from step 1 to construct the regressors $q\pi$ and $q\pi(1-\pi)$, and estimate the other parameters by GLS.

Next, we test the theoretical prediction that the risk premium enters all equations with the same coefficient, that is, $\gamma_j = \gamma$ for every currency j. We exploit the fact that Wald test statistics obtained using the DOLS and DSUR estimators are asymptotically chi-square under the null hypothesis (Stock and Watson 1993, Mark *et al.* 2005) that the parameter γ is the same across exchange rates. For both the DOLS and DSUR estimates, the Wald test cannot reject the null hypothesis of our theoretical cross-equation restriction at any conventional significance level. Finally, we re-estimate the DSUR system imposing this restriction on γ. Tables 14 and 15 present the results.

The estimates of β, γ and q_0 are in line with the results obtained so far, and thus corroborate our findings. Moreover, for all countries except Norway, Australia and New Zealand, the

TABLE 14 Dynamic SUR (DSUR) estimates with cross-equation restriction on γ.

	EUR	USA	JAP	CHE	CAN
$q\pi$	0.212***	0.194***	0.346***	0.235***	0.212***
	(0.038)	(0.027)	(0.057)	(0.039)	(0.047)
$q\pi(1-\pi)$	0.377***	0.377***	0.377***	0.377***	0.377***
	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)
$i^* - i$	2.712***	3.384**	−8.282***	1.288	12.965***
	(0.717)	(1.437)	(1.720)	(1.176)	(1.17)
Constant	−0.337***	−0.444***	−5.348***	−0.399***	−0.716***
	(0.007)	(0.005)	(0.014)	(0.019)	(0.007)
q	−0.037	−0.059**	−0.006	−0.081***	−0.078***
	(0.025)	(0.023)	(0.030)	(0.025)	(0.027)
q_0	0.273	0.273	0.273	0.273	0.273

Notes: The table reports DSUR estimates (and standard errors in parentheses) for the bilateral log exchange rates of the BP against the euro (EUR), the US dollar (USA), the Japanese yen (JAP), the Swiss franc (CHE) and the Canadian dollar (CAN). Here, π is the Leave probability and q is the probability of the referendum being held, taking value q_0 before the Referendum Act and 1 afterwards. The bottom part of the table reports the estimate of q_0. Also, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates. The value of q_0 and the long-run error covariance matrix are estimated in the first step by DOLS, imposing the equality of q_0 across currencies. The long-run covariance matrix is estimated using a Bartlett kernel with Newey–West fixed bandwidth 6. The second step is estimated by GLS. The coefficient of $q\pi(1-\pi)$, i.e. γ, is constrained to be the same across currencies. The sample goes from 27 May 2015 to 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

TABLE 15 More dynamic SUR (DSUR) estimates with cross-equation restriction on γ.

	DAN	SWE	NOR	AUS	NZL
$q\pi$	0.215***	0.200***	0.209***	0.251***	0.292***
	(0.037)	(0.040)	(0.052)	(0.055)	(0.050)
$q\pi(1-\pi)$	0.377***	0.377***	0.377***	0.377***	0.377***
	(0.075)	(0.075)	(0.075)	(0.075)	(0.075)
$i^* - i$	3.169***	1.298	10.361***	5.554***	4.461***
	(0.461)	(1.166)	(1.208)	(1.320)	(1.097)
Constant	−2.330***	−2.58***	−2.613***	−0.869***	−0.995***
	(0.007)	(0.013)	(0.010)	(0.025)	(0.031)
q	−0.045*	−0.027	−0.060**	−0.052*	−0.018
	(0.025)	(0.026)	(0.029)	(0.030)	(0.029)
q_0	0.273	0.273	0.273	0.273	0.273

Notes: The table reports DSUR estimates (and standard errors in parentheses) for the bilateral log exchange rates of the BP against the Danish krone (DAN), the Swedish krone (SWE), the Norwegian krone (NOR), the Australian dollar (AUS) and the New Zealand dollar (NZL). Here, π is the Leave probability and q is the probability of the referendum being held, taking value q_0 before the Referendum Act and 1 afterwards. The bottom part of the table reports the estimate of q_0. Also, $i^* - i$ is the difference between the foreign and domestic (UK) interest rates. The value of q_0 and the long-run error covariance matrix are estimated in the first step by DOLS, imposing the equality of q_0 across currencies. The long-run covariance matrix is estimated using a Bartlett kernel with Newey–West fixed bandwidth 6. The second step is estimated by GLS. The coefficient of $q\pi(1-\pi)$, i.e. γ, is constrained to be the same across currencies. The sample goes from 27 May 2015 to 30 June 2016.
***, **, * indicate significance levels 1%, 5%, 10%, respectively.

point estimates of θ are now closer to the UIP-implied value 1, compared to the single-equation estimates.

6 | CONCLUSIONS

In this paper, we analyse the role that political uncertainty plays in currency markets. We choose not to follow the prevalent literature, which typically starts by building indexes based either on financial volatility measures or on the frequency of political news featuring particular keywords. Instead, we choose a very specific political event, Brexit, relative to which we can define and measure the probability distribution of the event under scrutiny. We do so by exploiting data from bookmakers' odds.

Assuming that, during the period of analysis, the most relevant source of political uncertainty for the BP is the Brexit Referendum outcome, our approach allows us to pin down the conditional exchange rate expectations and to exploit the event's binary nature, Leave or Remain, for deriving the 'correct' measure of uncertainty, the second moment of the distribution. This measure provides us with a theoretically 'correct' definition of time-varying risk premium associated with currency markets, and is robust to the inclusion of popular measures of political uncertainty and to time-varying risk aversion.

Our simple model of portfolio choice predicts that a rise in the Leave probability should (likely) be associated with a depreciation of the British pound (BP) relative to other currencies, and that overshooting should occur as investors move away from a 'riskier' BP due to the volatility effect.

According to the standard presumption, investors' risk aversion rises when a bad state of the world, the vote for Brexit, materializes, and this should imply a flight out of the BP. Our model implies exactly the opposite: following the result of the vote, the uncertainty about the referendum outcome vanishes, and this tends to strengthen the BP. The empirical evidence confirms our theoretical result, and suggests an intuitive explanation for the BP (2%) appreciation following Boris Johnson's electoral victory in December 2019. Given the prime minister's commitment to Brexit, this could be explained only by a reduction of the uncertainty surrounding Brexit.

Our results are also important in the context of the so-called 'uncovered interest parity anomaly': by relying on bookmakers' odds for deriving expectations and political risk premia, and by allowing for cross-equation restrictions derived from portfolio choice, we go quite some way to solving the puzzle of the 'wrong' sign of the interest rate differential.

Finally, from a methodological point of view, we relate uncertainty and structural breaks. In particular, the resolution of uncertainty concerning the referendum is shown to generate structural breaks of the equation parameters, since these parameters are not 'structural' but reflect the way in which market participants update their expectations following the arrival of new information. Thus our approach seems suitable for a variety of fruitful applications to modelling the relationship between political uncertainty, expectations, asset prices and structural breaks.

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ENDNOTES

[1] For example, a study by HM Treasury (2016) concluded that 'if we take as a central assumption that the UK would seek a negotiated bilateral agreement, like Canada has, […] our GDP would be 6.2% lower, families would be £4300 worse off and our tax receipts would face an annual 36 billion black hole'.

[2] See, for example, Taylor and Guarascio (2016).

[3] Mike Smithson, founder and editor of PoliticalBetting.com, defined Brexit as 'the biggest political betting event of all time, anywhere'. According to Collinson (2016), 'More than £40m was gambled in the biggest political betting event in British history'.

[4] We focus on ten currencies, which are representative of the major British trade partners: the euro, the US dollar, the Japanese yen, the Swiss franc, the Canadian dollar, the Danish krone, the Swedish krona, the Norwegian krone, the Australian dollar and the New Zealand dollar.

[5] This has also been documented in recent papers that focus on exchange rates and Brexit (e.g. Korus and Celebi 2018; Hanke *et al.* 2018; Auld and Linton 2019; Clark and Amen 2017). None of these papers considers second moments—so they are unrelated to political uncertainty—or builds a model to interpret the relation between politics and exchange rates 'structurally'.

[6] These include the exchange rate 'disconnect' from macro-variables, the empirical failure of the UIP condition, the persistence of real exchange rate deviations from purchasing power parity, and lack of international risk-sharing.

[7] Throughout the analysis, we abstract from default risk. This seems a reasonable first-order approximation, given that we are considering countries where default risk is considered to be fairly small—both in the sovereign credit default swap market and according to all major credit rating agencies.

[8] The term ω^2 follows from the property that for a random variable X, we have $\mathrm{Var}(\omega X) = \omega^2\ \mathrm{Var}(X)$.

[9] This assumption is relatively innocuous in the short run, when the supply of currencies hardly moves.

[10] In the Appendix, we show that the portfolio variance falls after a Leave vote, as in the base model, provided that the probability of a hard Brexit, prior to the referendum, is sufficiently large. In other words, beliefs before the vote should be that a Leave outcome would likely lead to serious consequences as in the 'hard' scenario, rather than it being nullified by subsequent treaties, which is plausible given the heated debate at the time (e.g. Sampson *et al.* 2016).

[11] For simplicity, throughout the paper we refer to 'countries', treating the euro area as a single country.

[12] Probabilities are calculated from odds using the standard definition that the odds equal the ratio of favourable to unfavourable outcomes. For example, odds of 1 : 9 characterize an event that has one positive and nine negative realizations every ten, so that the probability of the positive realization is $\pi = odds/(1 + odds) = 1/10$. These odds are derived from the transactions placed on the bets market, thus they aggregate risk attitudes of participants. We thus assume that risk attitudes in the currency and betting markets do not differ significantly.

[13] Investors in PredictIt buy assets whose price represents the probability of a certain outcome (e.g. Leave victory) and can hold the asset until maturity (the referendum) or trade it before maturity, at the ongoing price. In order to comply with US regulators, PredictIt caps the size of trading positions (see Wolfers and Zitzewitz 2018). We do not have information on individual traders; however, PredictIt described to us its investors as follows: 'they are affluent (100–200K annual income), well-educated millennials, aged 22–35, living in metropolitan areas like NYC, DC, Philadelphia, Dallas, Chicago and San Francisco. Most traders work in finance, law, politics, and technical fields, such as mathematics, statistics and economics. They are politically diverse, with Democrats, Republicans, libertarians, 'unaffiliated' or 'no party' affiliates all using the site. There are about 30–35K active traders on PredictIt (defined as someone with money in their account) at any given time, with people entering and exiting markets regularly. Nearly 180,000 people have opened an account at time of writing.'

[14] See, for example, Holehouse and Hughes (2015) and Wright and Grice (2015).

[15] See, for example, Sparrow (2016).

[16] Estimates of an error correction model—available on request—confirm the results presented here and show that the exchange rates converge over time to their equilibrium relationship.

[17] Specifically, we use the Gauss–Newton algorithm for the numerical minimization of the sum of squared residuals. For DOLS and OLS, we optimize over all parameters simultaneously. To ensure that we detect a global minimum, we repeat the optimization 1000 times, each time using random draws from uniform$(-10, +10)$ distributions as starting values. For FMOLS, estimating q_0 and the other parameters simultaneously is unfeasible. Thus we follow a two-step procedure in each iteration of the optimization algorithm: first we set a value for q_0, then we estimate the remaining parameters using conventional (linear) FMOLS, conditional on q_0. Thus in this case, we optimize numerically over q_0 only. For DOLS and OLS, the standard errors of the parameters are obtained using the Newey–West heteroscedasticity and autocorrelation (HAC) estimator. For FMOLS, the standard error of q_0 is calculated as $\widehat{\sigma}_{lr}(\tilde{h}/2)^{-1/2}$, where $\widehat{\sigma}_{lr}$ is the long-run standard deviation of FMOLS residuals, and $\tilde{h}$ is the second derivative of the objective function (the sum of squared residuals) with respect to q_0. This approach follows conventional methods for calculating standard errors of non-linear regressions from the Hessian matrix (e.g. Amemiya 1983). For the other parameters, we calculate ordinary FMOLS standard errors, conditional on q_0.

[18] Note that the standard errors of q_0 in Table 2 are obtained without imposing that q_0 lies between 0 and 1. So, for instance, the 90% confidence interval of the DOLS estimate for ROW^F has a lower bound of -0.03 and an upper bound of 0.53, approximately. To impose $0 < q_0 < 1$, we can use the logit transformation of q_0, denoted by λ_q; i.e. we can replace q_0 with $\exp(\lambda_q)/(1 + \exp(\lambda_q))$ in the regression. In this case, while the point estimate of q_0 remains the same, the lower and upper bounds of the 90% interval are 0.07 and 0.60, respectively.

[19] The Wald test for the hypothesis $q_0\pi = 0$ has p-value 15%; for $q_0\pi(1 - \pi) = 0$, the p-value is 18%.

[20] More specifically, we use the risk reversal measured on options with a delta of 0.25 and a 3-month horizon.

[21] The data source is Refinitiv.

[22] We use high-frequency exchange rate data and option-implied volatility data for the US dollar, the euro, the Japanese yen and the Swiss franc, accounting for about 90% of the currency basket of the BP effective exchange rate.

[23] Thus, in this case, multilateral exchange rates and interest rates are calculated as weighted averages of the corresponding variables for the USA, the euro area and Japan only. Benetrix *et al.* (2019) provide only aggregate data on exposures to currencies other than the US dollar, euro, pound sterling, Japanese yen and renminbi. However, in 2015, the US dollar, the euro and the Japanese yen accounted for 84% of the UK's external assets and 92% of liabilities.

[24] The long-run covariance matrix is estimated using a Bartlett kernel with Newey–West fixed bandwidth 6.

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SUPPORTING INFORMATION

Additional supporting information can be found online in the Supporting Information section at the end of this article.

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APPENDIX. POST-REFERENDUM UNCERTAINTY

In this appendix, we extend the analysis of our baseline model to the case in which there is residual uncertainty in the post-referendum period, for instance, between a hard Brexit (HB) or a soft Brexit (SB). To keep things simple, we restrict attention to the case in which the event of a soft Brexit coincides with the outcome of the referendum being Remain—that is, the soft Brexit scenario is one in which subsequent agreements effectively 'undo' the impact of Leave on the British economy.

In this case, we have two possible outcomes: if Leave wins and a hard Brexit occurs, then the exchange rate is expected to be $e_H := \mathbb{E}(e'|\text{Leave}, \text{HB})$; otherwise, it is expected to be $e_L := \mathbb{E}(e'|\text{Leave}, \text{SB}) = \mathbb{E}(e'|\text{Remain})$. In this appendix, for simplicity, we set $q = 1$, which allows us to write $\mathbb{E}(e'|\text{Leave}) = he_H + (1-h)e_L$, where h is the probability that a hard Brexit occurs after Leave wins the referendum. Thus, prior to the referendum taking place, we have $\mathbb{E}(e') = \pi\ \mathbb{E}(e'|\text{Leave}) + (1-\pi)e_L = \pi he_H + (1-\pi h)e_L$.

Consider now the variance. After the referendum takes place, the variance is zero upon Remain, which we write as $\text{Var}(e'|\text{Remain}) = 0$. Moreover,

$$\begin{aligned}\text{Var}(e'|\text{Leave}) &= h\left[e_H - \mathbb{E}(e'|\text{Leave})\right]^2 + (1-h)\left[e_L - \mathbb{E}(e'|\text{Leave})\right]^2 \\ &= (e_H - e_L)^2 h(1-h).\end{aligned}$$

Moving one step backwards, to the pre-referendum period, we have that

$$\begin{aligned}\text{Var}(e') &= \pi h\left[e_H - \mathbb{E}(e')\right]^2 + (1-\pi h)\left[e_L - \mathbb{E}(e')\right]^2 \\ &= \pi h\left[e_H - \pi he_H + (1-\pi h)e_L\right]^2 + (1-\pi h)\left[e_L - \pi he_H + (1-\pi h)e_L\right]^2 \\ &= (e_H - e_L)^2 \pi h(1-\pi h).\end{aligned}$$

For the results in our model to reconcile the empirical evidence, one needs that the *ex ante* uncertainty before the referendum outcome is known is larger than the *ex post* uncertainty conditional on the Leave result, $\text{Var}(e') > \text{Var}(e'|\text{Leave})$, or equivalently, $h > 1/(1+\pi)$. Given that π has been around 30% on average in our sample, it follows that the probability of a hard Brexit after the Leave vote should be larger than $1/1.3 = 0.77$. Thus if the market's implied probability that a Leave vote would indeed result in a hard Brexit was greater than 77%, then our baseline results go unchanged. Note that this lower bound depends on the assumption that a soft Brexit is similar to a no Brexit scenario. If we assumed that a soft Brexit is somewhere in between the two, the lower bound would fall further.