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NOMINAL EXCHANGE RATE DYNAMICS AND MONETARY POLICY: UNCOVERED INTEREST RATE PARITY AND PURCHASING POWER PARITY REVISITED

Yossi Saadon and Nathan Sussman

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Centre for Economic Policy Research
33 Great Sutton Street, London EC1V 0DX, UK
Tel: +44 (0)20 7183 8801
www.cepr.org

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JEL Classification: F3, F31, F41, G15, E52

Keywords: purchasing power parity, uncovered interest rate parity, Exchange Rates, monetary policy, Inflation expectations, Balance sheet effects

Yossi Saadon - yosis@boi.org.il
Bank of Israel

Nathan Sussman - nathan.sussman@graduateinstitute.ch
Graduate Institute Geneva and CEPR

Nominal Exchange Rate Dynamics and Monetary Policy: Uncovered Interest Rate Parity and Purchasing Power Parity Revisited

Yossi Saadon and Nathan Sussman*

Abstract

The increasing globalization of trade in goods and services and the deepening of financial markets have reduced frictions that may impede the PPP and UIP relationships' operation in the short run. In this paper, we estimate the short term relative PPP and UIP relationships. Using data from Israel, which has a deep market for inflation expectations for 12 months, we show that relative PPP and UIP cannot be rejected. Deviations from equilibrium last less than a year. Data from Israel's capital account of the balance of payments shows that the deviations are not destabilizing. Our findings suggest that greater globalization, financial deepening, and developing a market for short-term inflation expectations contribute to monetary policy effectiveness.

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* Yossi Saadon, the Bank of Israel, Nathan Sussman, the Graduate Institute Geneva and CEPR. The views expressed in this paper do not necessarily represent the views of the Bank of Israel.

Introduction

Monetary policy's effectiveness and impact have been debated since the onset of the global financial crisis. A recent summary by Forbes (2018) emphasizes the increasing impact of global factors on domestic inflation, the Phillips curve, and monetary policy effectiveness, especially for advanced small open economies. Central banks have become increasingly concerned with developments in their currency's exchange rates, and some intervene in foreign exchange (FX) markets to defend their currencies against appreciation. Theoretically, the standard Mundell-Fleming model assumes that in a floating-exchange-rate regime, the domestic central bank can control monetary aggregates, which is more conducive to achieving inflation targets and output stabilization than under a fixed exchange rate. However, recent research has demonstrated that various financial frictions and the dominance of the US dollar and US financial markets may undermine monetary policy independence even with flexible exchange rates (Rey, 2016). Recent experience from the COVID-19 pandemic shows that capital outflows from emerging markets triggered exchange rate depreciation and increased local currency bond yields that undermined bond domestic monetary policy response (Hofmann et al., 2020).

One salient deviation from the equilibrium conditions underpinning the Mundell-Fleming model is the prevalence of carry-trading, which is considered one of the most profitable strategies in foreign currency trading and constitutes a significant factor in the foreign exchange markets of small open economies. Recently, Plantin and Shin (2018) modeled the potentially destabilizing effects of carry-trades. Carry trading may increase the deviations from the uncovered interest rate (UIP) relationship.

Purchasing power parity (PPP) and UIP equilibrium conditions, which underpin the standard Mundell-Fleming model, are appealing because they are based on a fundamental assumption—the absence of arbitrage—hence their popularity in textbooks. However, they require hard to fulfill conditions such as perfect competition and, for the UIP condition, deep financial markets with free capital flows. To *jointly* test the dynamics of nominal exchange rates related to the current and capital account, we also require flexible exchange rates. Moreover, since these conditions are forward-looking, short term market-based inflation expectations are required. These conditions limit the cases that can be used to examine the validity of these textbook economic conditions. Therefore, the extent to which market frictions undermine the UIP and PPP operation and hence the effectiveness of monetary policy is ultimately an empirical question.

In this paper, we test whether the UIP and PPP conditions hold in the short run in a market we believe matches more closely the assumptions behind these no-arbitrage conditions. Our contribution is to use data from Israel, an advanced small open economy that can be considered a price taker in global consumer goods and financial markets. We chose Israel because it is unique: short-term inflation-indexed bonds are traded in a relatively deep market because of its high inflation history. This allows us to use *forward-looking* 12 months' market-based breakeven-inflation rates that are publicly available from the Bank of Israel. Unfortunately, data on market-based inflation expectations for 12 months for additional advanced small open economies are either unavailable or available for only a short sample period (Adeney, 2017). Because of its inflation history, data from the Israeli economy also offers greater variation in the regressors, allowing to recover the UIP or PPP coefficients with higher precision. Moreover, the Israeli economy is an interesting test case for several other reasons: it is a small open economy that transitioned from an emerging market to an advanced economy; significant changes occurred in the foreign exchange market; the disinflation process came to fruition during the sample period; there was significant intervention in the foreign exchange market during part of the period; and as in most of the world's economies, the consequences of the financial crisis were dealt with an accommodative monetary policy that lowered interest rates to almost zero. Thus, we subject our estimation to a wide variety of changes in the environment in which exchange rates are determined.

Following a brief theoretical exposition of the joint determination of the UIP and PPP relationship in a small open economy, we estimated the relationship between the percentage change in the Israeli New Shekel (NIS) and the US Dollar over twelve months as a function of the twelve months' inflation expectations differential between Israel and the US prevailing at the beginning of the period. We could not reject the hypothesis that relative expected PPP holds and that the coefficient of the expected inflation rate differential is equal, as in theory, to 1. Estimating a vector error correction equation for the relative PPP relationship, we find that following a shock, it takes on average a year for the exchange rate to revert to its equilibrium level.

(Figure 1, about here)

The second step was to estimate the UIP relationship: We regressed the yield difference between BOI bonds and US treasuries on the expected change in the exchange rate. A variety of specifications shows that UIP holds with a coefficient of 1. We obtained these results by estimating a simple OLS regression of the interest rate differential on the forward change in the exchange rate. Finally, we estimated the joint equilibrium conditions using two-stage least squares estimation, where the PPP relationship is the first stage, and the UIP relationship is the second stage. Again, we obtained a coefficient of 1. This estimation (Figure 1) shows that the UIP relationship anchors the dynamics of the exchange rate. A vector error correction estimation for the UIP relationship shows, as can be gleaned from Figure 1 that the exchange rate reverts to the equilibrium relationship within a year.

Our results were obtained using the simplest possible specifications, and we found none of the puzzles reported in the literature (Engel, 2016). Our aim in this paper is not to fully account for the deviations of exchange rates from the PPP or UIP relationships nor to forecast exchange rates. We show that we cannot reject PPP or UIP when necessary conditions for their existence are more closely met. Namely: price-taking behavior, international goods, and financial markets integration, and the existence of a market for inflation expectations for the short run. These results are obtained for a wide variety of economic and policy changes, including a period when interest rates in Israel were at the effective lower bound. The deviations from UIP are short-lived, and we do not observe sustained deviations. Therefore, remaining frictions do not offer carry-trade opportunities beyond the very short-run. Owing to the relatively rapid corrections, they are highly risky.

Using balance of payments data, we can reject the hypothesis that deviations from UIP or PPP affect net portfolio investment. We also show that some of the deviations were probably induced by monetary policy. These findings are also corroborated when we subject the data from Israel to the test proposed by Hofmann et al. (2020). Unlike the findings from most small open economies and emerging markets, the US Dollar return of five-year domestic currency bonds does not differ from the NIS return on these bonds.

We cannot reject the relative expected PPP relationship. However, this does not imply that the exchange rate pass-through to the economy is high. It only implies that the *expected* pass-through is high. Obviously, as shown in Figure 1, exchange rates are subject to high, unexpected volatility. The pass-through from actual changes in the exchange rate will likely decrease with this volatility (Devereux, Engel, 2002). This finding is also consistent with (Forbes, Hjortsoe, and Neova, 2018), who show that exchange rate shocks caused by monetary policy shocks have a greater impact on prices. Since according to our framework, monetary policy shocks affect inflation expectations.

The main policy implication of our findings from Israel for small open economies and emerging markets is that developing a domestic currency bond market and having a flexible exchange rate are necessary but not sufficient conditions to allow for effective monetary policy and mitigation of deviations from UIP. It is also necessary to develop a deep market for hedging inflation, preferably with inflation-linked bonds.[1] We believe that in a world with rising goods', services' and financial integration and the availability of forward-looking inflation hedging assets in small open economies, the results we obtained for Israel will become commonplace elsewhere. Recently, Cavallo et al. (2018) showed using data from online platforms that PPP holds quite well for these products. This provides a more positive outlook for monetary policy in small open economies with flexible exchange rates and inflation targeting. It also opens up another channel of increasing monetary policy effectiveness, namely ensuring the continuation of trade liberalization of goods and services and the development of financial markets.

The paper is structured as follows. We begin with a summary of relevant literature. In Section 2, we provide the theoretical model we estimate. In Section 3, we estimate the UIP and PPP relationships and describe the data used. In Section 4, we provide some extensions and robustness checks. We conclude with a summary of the results and policy implications.

1. Related literature

Our paper is closely related to the extensive literature that estimates the UIP relationship. The extensive literature does not support, for the most part, the existence of the UIP relationship, especially for the short run.[2] Moreover, many studies show a negative interest rate gap coefficient for short-term interest rates consistent with the carry trade literature.[3] Froot and Thaler (1990) reviewed dozens of empirical studies that tested the relationship and found that the average coefficient was -0.88. Bansal (1997) showed that the failure of UIP is more significant for industrialized economies compared with developing economies. Bansal and Dahlquist (2000) and Ballie and Kalic (2006) argue that the existence of the relationship depends on whether the interest rate gap is positive or negative. A deviation from the UIP relationship will appear when the US interest rate is higher than the local interest rate.

Chinn and Meredith (2004) assessed the UIP relationship for 5 and 10 years for the G7 countries. They found stronger support for the existence of the relationship for long-term interest rates than short-term interest rates of 12 and three months. Their study showed that the

[1] Inflation swaps are not traded in a liquid market and the quotes are generally not available to the general public.
[2] For example, Bekaert and Hodrick (1993), Engel (1996), Mark and Wu (1998), and Weber (2011), and others. On the other hand, a few articles do find support for this relationship: Flood and Rose (1996), Bekaert and Hodrick (2001), Baillie and Bollerslev (2000), Chaboud and Wright (2005), Bayaert, Garcia-Solanes and Perez-Castejon (2007), Omer et al. (2013) and Tang (2011).
[3] For example, MacDonald and Taylor (1992), Isard (1996), McCallum (1994), Engel (1996), and Chinn and Meredith (2004).

interest rate spread coefficient was positive and closer to 1 than to 0. In another and more recent article, Chinn and Quayyum (2012) again supported the conclusion that the UIP relationship is valid for long-term interest rates. Coffey, Hrung, and Sarkar (2009), who refer to the 2008 financial crisis, assert that the basic assumption of UIP that the financial markets function effectively enough to prevent arbitrage was invalid in the 2008 financial crisis. Investors during this period, therefore, preferred to wait in currency positions until the crisis subsided.

In a recent comprehensive study, Engel (2016) reviews articles addressing UIP and discusses the apparent contradiction between the UIP theory and the empirical findings. According to these studies, the exchange rate of countries with a high-interest rate tends to be revalued more than according to the interest rate spreads based on the UIP model. As found in other studies, he also found that the real exchange rate converges to the UIP relationship over time. Engel tested the UIP on six leading economies (Canada, France, Germany, Italy, Japan, and the UK) in 1979–2009 compared to the US. His study showed that the exchange rate risk plays an important role in explaining this contradiction. He includes liquidity of assets as a factor that can explain his findings with respect to the UIP relationship. Our contribution is to test the UIP for a small open economy rather than for large ones that may be considered market makers. We also use short term inflation expectations obtained from financial markets rather than using inflation expectations based on lagged inflation. Finally, while we follow Engel and emphasize the estimation of these relationships for the short run, we use annual observations that correspond to the maturity of the financial assets that we use, rather than convert them into monthly maturities by taking the 12th root.[4]

Our paper is also related to the literature on carry-trades and the recent emphasis on exchange rate related balance sheet effects of monetary policy instead of the more traditional expenditure switching cannel (Plantin and Shin, 2018). Suppose that the central bank of a small open economy raises interest rates to cool down the economy. According to the Mundell-Fleming framework, this will result in capital inflows and appreciate the currency. However, carry traders will lend to the high-interest rate economy. The increasing capital inflows will counteract monetary policy by stimulating domestic investment and consumption through a balance sheet effect. This exposes the small economy to financial instability and loss of control over inflation. While this strategy is exposed to considerable risk (Doskov and Swinkels (2015) and Burnside et al. (2007)), it is often pursued when there are deviations from uncovered interest rate parity (UIP). As a result, they make the deviations from UIP greater and longer-

[4] In the Robustness section we follow Engel (2016) and show our results also hold, but with lower significance.

lasting (Baillie and Chang, 2011).[5] Thus, under conditions outlined by Plantin and Shin (2018), carry trading may increase the deviations from the UIP.

The theoretical and empirical analysis of carry-trading as a destabilizing activity that undermines the effectiveness of monetary policy rests on the existence of various frictions in markets and inappropriate monetary policy. For example, in Plantin and Shin (2018), the destabilizing outcome is partly the result of sticky non-traded goods prices that undermine the PPP relationship and partly the result of passive monetary policy that underestimates the balance sheet effects of capital flows. These theoretical considerations fit well with the empirical literature on the UIP. Our contribution is to show that when some of the frictions are removed, we do not find sustained deviations from UIP. This suggests that the carry trade under such circumstances is riskier.

Our paper also relates to the extensive literature on purchasing power parity (PPP). Previous research was summarized by Taylor and Taylor (2004). They reviewed the extensive literature on the PPP and concluded that the PPP relationship could not be rejected in the long run. More recent research has emphasized frictions in goods markets that explain the deviations from PPP in the short run. One example is Sarno et al. (2004), who model deviations from PPP that occur because of non-linear frictions such as transaction costs. Another version of price stickiness that differs by sector is found in Carvalho and Nechio (2011). Recently, Engel (2018) and Engel and Zhu (2018) revisited the PPP puzzle (Rogoff (1996, p. 647-648)) that argues that the real exchange rate converges very slowly (half-life of 3 to 5 years), much slower than nominal prices and therefore contradicts PPP. Engel (2018) offers a New-Keynesian model with Calvo pricing that is centered around sticky prices. At the same time, Cavallo et al. (2018) use micro-level data from online shopping platforms and show strong evidence for PPP. Their results suggest that there are empirical issues associated with price level statistics collected by national statistical bureaus. Our results support Cavallo et al. (2018), and at the same time, we find no puzzle in the convergence of Israel's real exchange rate with the US.

2. Theoretical Framework

Our point of departure is the existence in the long run of two classic behavioral equations: the first is the *relative* purchasing power parity (PPP) equation, which states that the change in the exchange rate is equal to the difference between the inflation rates in two economies. The

[5] In practice, profits from carry trading are usually positive, although it has caused heavy losses over short periods in the past (see Doskov and Swinkels, 2015). However, Fama (1984), Chinn & Meredith (2004), and Frenkel & Poonawala (2010), like many others, show that the profit and loss depend on the, country, and time period.

second is the uncovered interest rates parity (UIP) equation. The difference between interest rates of similar assets in two economies is equal to the expected depreciation of the domestic currency against the other economy's currency. Most studies UIP that used data from industrialized countries also examined the unbiasedness hypothesis that assumes risk-neutral individuals and a zero or constant risk premium.

Under the rational expectations assumption, the forward-looking relative PPP equation for estimation can be written as:

$$(1) \quad (Es_{t+1} - s_t) = \alpha_P + \beta_P(E\pi_{t+1} - E\pi^*_{t+1}) + u_{P,t+1}$$

where ***s*** represents the log of the exchange rate in terms of units of domestic currency to a unit of foreign currency; $E\pi_{t+1}$ represents the expected inflation rate and $E\pi^*_{t+1}$the expected inflation rate abroad. The subscript *P* denotes the PPP equation. The null hypothesis is that $\beta_p = 1$.

The UIP condition is commonly tested using Fama's (1984) regression:

$$(2) \qquad (Es_{t+1} - s_t) = \alpha_U + \beta_U(i_t - i^*_t) + u_{U,t+1}$$

where *i* represents the domestic nominal interest rate; $i*$ represents the nominal interest rate abroad and α represents the constant risk premium. The subscript *U* denotes the UIP equation. Under UIP, the null is that $\alpha_u = 0 \; and \; \beta_u = 1$.

Following Engel (2016), we define the real exchange rate q_t as:

$$(3) \; q_t \stackrel{\text{def}}{=} s_t + \log(p^*_t) - \log(p_t \;)$$

where *p* represents the domestic and p* the foreign price level (CPI).

The real ex-ante interest rate *r* is defined as:

$$(4) \quad r_t = i_t - E\pi_{t+1}$$

3. Estimation of the PPP and UIP—the case of Israel.

In this section, we present the data and the basic estimation results that support our hypothesis that the PPP and UIP conditions cannot be rejected for Israel's case. Our sample period of monthly observations begins in January 1996 and ends in July 2018, the latest point in time available. During this period, there were many changes in the economic environment and policies in Israel that make the study of the Israeli case interesting and, at the same time, challenging for estimating the short-run PPP and UIP relationships.

During this period, there were three global financial crises: LTCM in 1998, the 2000 dot.com crisis, and the global financial crisis (GFC) that began in 2008, as well as a domestic crisis of the Second Intifada (2000-2004). During this period, there were structural changes and changes in Israeli economic policy. The most prominent of them were the disinflation process that ended in 2003 and the change in the Bank of Israel's policy in the foreign exchange market in 2008. The exchange rate regime witnessed several developments: at the beginning of our sample—from 1996 to the end of 1997—the exchange rate operated within a band of +/- 7%. After that, the band's width was increased to 60%, and the Bank of Israel stopped intervening in the FX market. In 2005, the exchange rate band was officially removed. In 2008, as part of unconventional monetary policy, the Bank of Israel started to intervene in the FX market, a policy that is still carried out at the time of writing this paper. Finally, in September 2010, Israel became a member of the OECD, and in that year, Israel also exited the MSCI emerging market index and moved to the developed markets index.

3.1. Data for estimation

Our key contribution is to use forward-looking inflation expectations in the empirical estimation of the PPP equation (1). For Israel, we use the inflation expectations for 12 months estimated from the market for inflation-indexed bonds and the 12-month Bank of Israel nominal bond. The data are publically available from the Bank of Israel website. For the US, we used the Michigan survey of consumers provided by the Federal Reserve Bank of St. Louis (FRED). Our decision to focus on Israel is driven by either the absence of market-based inflation expectations or their very short sample size in small advanced open economies (Adeney et al., 2017).

Many studies estimate inflation expectations from actual inflation or lagged inflation. Recently Engel (2016) used a vector error correction model to extract inflation expectations and calculate the ex-ante interest rate. The recourse to estimation is partly because a series of market-based one year ahead inflation expectations (either from breakeven inflation or swaps) does not exist or does not have a long enough history for many countries. The advantage of using data from Israel is that we use available assets that are common knowledge and have existed for a long time (a legacy of the high inflation era of the 1980s), rather than those estimated ex-posts by researchers. As shown in Figure 2, there is quite a substantial difference between lagged or actual inflation and inflation expectations. As we will show later, using expected inflation differentials rather than actual or lagged inflation makes a significant difference in estimating the PPP relationship.

(Figure 2, about here)

The interest rate differential we focus on is the one-year nominal yield spread between a 12-month constant maturity Treasury bill (data from the Federal Reserve Bank of St. Louis (FRED)) and the yield on the one-year Bank of Israel bill—the Makam. As Figure 3 shows, although the yield on the US Treasury bill was almost unchanged following the onset of the crisis, that of Israel did. This provides a variation in the spread even when the US, and later Israel, were at the effective lower bound.

(Figure 3, about here)

As customary in the literature (Engel, 2016), we denote the expected depreciation by using the ex-post realized annual change in the USD-NIS (Israeli New Shekel) exchange rate.

We introduce two additional variables that will be used in our robustness checks. The first is Israel's country risk premium. We use the Bloomberg quote of the CDS on 5-year bonds to measure Israel's country risk. Unfortunately, the data is available only from July 2002. Because Israel was classified as an emerging market up to 2010, when it joined the OECD, we produced an estimate of Israel's CDS by using a forecast from a regression of Israel's CDS on the EMBI index for the period 2002 to 2007, when we have both measures. We detail this estimation in Appendix 1. Another variable we introduce for our robustness checks is the monthly amount of foreign exchange rate purchases made by the Bank of Israel.[6]

To assess the balance sheet effects of deviations from UIP or PPP, we use quarterly portfolio investment data, and other investment data from the balance of payments account available from the Bank of Israel.

3.2. Estimation and results

3.2.1. Estimating the PPP relationship

In this subsection, we estimate the forward-looking relative PPP: equation (1). Since inflation expectations and the expected change in the exchange rate are determined simultaneously, the estimation of equation (1) does not imply causality. Since we can reject the null of both individual and common unit-roots for the variables in our sample, we can view the estimation

[6] The data on monthly FX interventions is not publically available.

of equation (1) as the cointegrating equation.[7] We then estimate an error correction model to obtain the length of time it takes the nominal exchange rate to return to the PPP relationship.

The results show (Table 1) that the expected inflation differential coefficient is equal to 1, and we cannot reject the hypothesis that the intercept is zero. When we estimated a vector error correction model (VECM), we obtained that the speed of adjustment of the exchange rate to the PPP relationship is ten months.

In theory, both variables that make up the PPP equation are endogenous. However, using the expected inflation differential, rather than the actual inflation differential, we regress the ex-post change in the exchange rate on the ex-ante inflation differential. This should avoid issues of simultaneity. Nevertheless, we re-estimate the PPP equation (1) using instruments for the difference in expected inflation between Israel and the United States. Our instruments are the lagged actual inflation differential in the previous year, which is commonly used to estimate inflation expectations (Engel, 2016) and the lagged Fed policy rate—which under an inflation target regime is an indicator of the Fed's view on inflation. The results in the bottom panel of Table 1 show that our previous results hold. The two-stage least squares estimation enables us to test the exogeneity assumption of the expected change in the inflation differential. We find that though we can reject the weak instruments hypothesis, we cannot reject the hypothesis that the expected inflation differential is exogenous.

(table 1, here)

3.2.2. Estimating the UIP Relationship

In this subsection, we estimate the Fama regression (equation (2)). As in the PPP equation estimation, the equation does not imply causality. Since we reject the null of unit-root, we proceed to estimate the UIP equation as a cointegrating equation. As before, we then estimate the VECM to obtain the speed of adjustment to deviations from the UIP relationship. The results (Table 2) show that we cannot reject the null that the coefficient of the interest rate differential is 1 and that the intercept is equal to zero. The speed of adjustment of the exchange rate to deviations from the UIP (results of the VECM) turns out to be seven months. It is quite natural that the convergence of the exchange rate to financial market conditions (UIP) is faster than the convergence to the goods market equilibrium conditions (PPP).

[7] The unit root tests, detailed in Appendix 2, on the residuals of equation (1) confirms that the variables are cointegrated.

For similar considerations as in the PPP equation estimation, we estimate a two-stage least squares version of equation (2). Our instruments for the interest rate differential are Israel's country risk premium (captured by the CDS on 5-year bonds), and the Bank of Israel lagged policy rate. The results in the bottom panel of Table 2 show that our previous results hold. The two-stage least squares estimation allows us to test the exogeneity assumption of the interest rate differential. We find that though we can reject the weak instruments hypothesis, we cannot reject the hypothesis that the interest rate differential is exogenous.

(Table 2 about here)

3.2.3. Balance sheet effects of the impact of deviations from PPP and UIP.

In the previous subsections, we showed that relative expected PPP and UIP hold in Israel's case and that deviations from the equilibrium relationships tend to revert within less than a year. As Plantin and Shin (2018) argue, the deviations from UIP produce balance sheet effects that can potentially destabilize the economy (Rey, 2016). The findings we report above suggest that for the case of Israel, these are not major concerns. Mantzura and Shreiber (2016) recently documented the sharp increase of foreigners' holding of Makam bills in the wake of the opening interest rate differential with the US in 2010 and early 2011 (Figure 3). These were coupled with an appreciation of the currency and could be viewed as destabilizing manifestations of the carry trade. In this subsection, we test whether the deviations from the equilibrium UIP relationship impact the portfolio and other investments' account of the balance of payments.

Since our balance of payments data are quarterly, we first re-estimated equation 2 to obtain the quarterly residuals from UIP. We then estimated the following regressions of capital flows, $K_{i,t}$ on the residual from equation (2), i=1 for investments by nonresidents, i=2 for investments by residents, and i=3 for net flows.

$$(5)\;\; K_{i,t} = \alpha_{i,K} + \beta_{i,K} u_{U,t} + u_{K,t}$$

The results (Table 3) show that nonresidents invest more in Israel when the exchange rate deviates below what UIP implies, which can be viewed as destabilizing: when interest rates in Israel are high, and the exchange rate is appreciating (Plantin and Shin, 2018). On the other hand, residents react in the opposite direction and do not seem to coordinate with the destabilizing forces. Note that the reaction of nonresidents to deviations is much faster than that of residents. The interaction of both types of investors shows that the net investment flows are not affected by the deviations from UIP; we cannot reject the hypothesis that $\beta_{3,K} = 0$. Granger causality test confirms these results. In particular, we can reject the hypothesis that nonresidents cause, in the Granger sense, the behavior of residents.[8]

3.2.4. **Evidence for UIP from the five-year bond market**

Theoretically, the issuance of domestic-currency denominated bonds in emerging economies alleviates concerns of 'original sin' and allows monetary policy to be more effective. Recent research by Hofmann et al. (2000a, 2000b) compares the domestic-currency rate of return on domestic currency bonds with the US Dollar return on these bonds. They plot the two returns against the change in the bonds' yield (in percentage points). Theoretically, when UIP holds, there should be no difference between the rate of return in domestic and foreign currency. Otherwise, an arbitrage opportunity exists.

The data presented in Hofmann et al. (2000a, 2000b) shows that the credit risk of bonds fluctuates with the spot exchange rate for small open economies and emerging markets. A deprecation of the spot exchange rate is associated with lower returns on domestic bonds when measured in UD dollar terms than their return in domestic currency. This finding represents a deviation from the UIP and also from the covered interest rate parity (CIP).

We reproduced Hofmann et al. (2020b) ***figure 1*** for Israel. We used weekly data from January 2008 to July 2020. Our results show (Figure 4) Israel is unique: on average, the returns on five-year domestic currency bonds are ***identical*** to their US Dollar return. [9] This result provides additional evidence there are no systematic deviations from UIP in Israel. The credit risk of domestic currency bonds is insulated from spot exchange rate shocks. The results also confirm the absence of significant balance sheet effects we documented in the previous sub-section. This allows monetary policy to be much more effective than in emerging market economies and advanced small open economies.

(Figure 4 about here)

4. Extensions and robustness checks

Using the data for Israel and the US, we cannot reject the hypothesis that during the sample period from 1996 to 2017, the PPP and UIP relationships hold and that deviations from these

[8] See Table A2 for Granger causality results.
[9] The slopes of the regression line are -5.86 (s.e. 0.44) for Dollar returns and -6.12 (s.e. 0.13) for domestic currency return. Of course, the U.S Dollar returns are more volatile because the spot exchange rate is more volatile than the bond prices.

conditions converge back to equilibrium within less than a year. We obtained the results using the most straightforward specifications. We took advantage of the fact that for most of the sample period, Israel's monetary conditions differed from those in the US. Indeed, when monetary conditions are identical between two economies (interest rates and expected inflation), the PPP and UIP are *empirically* reduced to a random walk.

In this section, we provide some extensions and robustness tests. i) We test, within the framework of the PPP equation, whether the use of expected inflation differentials instead of actual (lagged) rates makes a difference in our estimations. ii) Since we found the PPP and UIP to hold, we subject our data to some of the puzzles estimated recently by Engel (2016). iii) The Bank of Israel intervened in the FC markets as part of the unconventional monetary policy following the GFC's onset. We will test whether FX interventions affect equilibrium conditions. iv) We test whether monetary policy can account for deviations from PPP and UIP. v) We re-estimate our PPP and UIP regressions using the monthly rate of change rather than annual. vi) We use bootstrap estimations to rule out that specific episodes determine our results.

4.1. Estimating the PPP relationship with actual inflation differentials.

In this subsection, we re-estimate the PPP relationship using 1. lagged inflation differentials as a proxy for expected inflation and 2. ex-post realized inflation and exchange rates.

The UIP conditions are essentially forward-looking, and since the expected change in the exchange rate is what matters. Since the PPP and UIP are two sides of the same exchange rate coin, we estimated the PPP using forward inflation differentials. Figure 2 plots the forward versus fully adaptive expectation differentials. The difference is noticeable. While the correlation between lagged inflation differential and forward inflation differential is quite high (0.8), when we substitute expected inflation differential with the lagged inflation differential, the coefficient drops from 1 to 0.55 (Table A1). This means that using lagged inflation, a Wald test confirms that we can reject the forward-looking PPP relationship.

What about the PPP relationship in real-time? Since the actual rate of change of the exchange rate and the actual inflation differential are determined simultaneously, we use 2SLS estimation to test for the 'real-time' PPP relationship. We instrument the inflation differential with the beginning of period lagged inflation differential, Fed, and Bank of Israel policy rates. We find (Table A1) that ex-post PPP also holds for Israel's case vis-a-vis the US.

4.2. Testing for PPP and UIP puzzles

Engel (2018) and Engel and Zhu (2018) address the PPP puzzle that shows that the real exchange rate converges to equilibrium too slowly than warranted by the PPP relationship. Specifically, the half-life of an AR(1) process of the real exchange rate is typically longer than three years (See also Sarno et al. , 2011). In contrast, the nominal exchange rate and relative prices of the two economies converge much faster. They find that the real exchange rate converges much faster for a fixed exchange rate than for flexible exchange rate economies.

Following Engel (2018), we test for the half-lives of the real exchange rate, nominal exchange rate, and relative price levels for Israel vis-a-vis the US. We also found that the point estimate of the real exchange rate's half-life is slightly longer (4 months) than that of the nominal exchange rate. However, the differences obtained from the AR(1) regressions are not significant at the 95% level (Table 4). As a robustness measure, we estimated rolling AR(1) regressions for two-year and seven-year windows and calculated the means of the coefficients of the AR(1) term. For the two-year rolling regressions, we cannot reject the hypothesis that the means are similar. Moreover, the mean of the rolling AR(1) coefficient on the real exchange is no longer the highest (bottom panel Table 4). Therefore, the evidence for a PPP puzzle is inconclusive.

(Table 4 about here)

Engel (2016) finds empirical evidence for two UIP puzzles that are contradictory in their implications. The first is the puzzle that relatively high interest rate countries have higher returns on short term deposits. The second puzzle is that countries with relatively high real interest rates have a stronger real exchange rate than would be implied by UIP.
Following the notation in Engel (2016), we define excess returns, ρ_{t+1}, as:

$$(6)\ \rho_{t+1} \stackrel{\text{def}}{=} {i_t}^* + s_{t+1} - s_t - i_t$$

The puzzle that Engel (2016) recently estimated is that excess return from investing in short term deposits abroad are positively correlated with higher real interest rates abroad:

$$(7)\ cov(\ E_t\rho_{t+1}, r_t^* - r_t) > 0$$

Using the definition of excess returns, ρ_{t+1}, from equation (6), Engel (2016) rewrites the Fama equation as:

$$(8)\ \rho_{t+1} = \zeta_s + \beta_s(i_t^* - i_t) + u_{s,t+1}$$

Our result that the UIP relationship holds suggests that we will be able to reject the first puzzle; we find (Table 5) that the simple correlation (equation (8)) is negative, but not significantly different than zero. Naturally, when we estimate the modified Fama regression, we find (Table 5) that we cannot reject the hypothesis that $\beta_s = 0$.

(Table 5, about here)

The second puzzle documented by Engel (2016) is formally stated as excess comovement of the stationary component of the exchange rate with respect to the ex-ante real interest rate differential. This means that the level of the exchange rate is strongly affected by deviations for UIP. The strong effect of the deviations on the exchange rate level implies that the higher interest rate (riskier) country has lower ex-ante risk premiums. Engel first regresses a model of the level of the real exchange rate (defined above, equation (3)) on the ex-ante real interest rate (from equation 4) differential:

$$(9)\quad q_t = \zeta_Q + \beta_Q(r_t^* - r_t) + u_{Q,t+1}$$

He finds that β_Q is positive, implying that when interest rates are higher abroad, the home exchange rate depreciates. This is not surprising. However, the puzzle emerges when he re-estimates (8), replacing the level of the real exchange rate as the dependent variable with the ex-ante sum of excess return premiums relative to their unconditional mean $E_t \sum_{j+0}^{\infty}(\rho_{t+j+1} - \bar{\rho}).$[10]

$$(10)\qquad E_t \sum_{j+0}^{\infty}(\rho_{t+j+1} - \bar{\rho}) = \zeta_P + \beta_P(r_t^* - r_t) + u_{p,t+1}$$

Engel finds that β_P is negative, which implies that higher interest rate currencies that are riskier have a lower ex-ante stream of risk premiums. When we estimate equation (8), we find (Table 4) that we observe the expected positive coefficient, which implies, according to UIP, that high *real* interest rate economies have an appreciated real exchange rate. However, when we estimate equation (9), we find, unsurprisingly, given our earlier results, no evidence for a puzzle (Table 4).

4.3. Accounting for the Bank of Israel FX intervention

As part of its unconventional monetary policy, the Bank of Israel purchased $US. Overall, FX purchases amount to more than $50 billion since 2008. The Median monthly intervention was $500 million, with the highest amount being $4 billion in August 2009 in the midst of the global financial crisis. In this subsection, we re-estimate equations (1) and (2) and control for FX purchases. We use two formulations: the first uses the monthly amount purchased and amount squared to allow for non-linearity in the effect of purchases. In the second, we use a dummy variable for months when the Bank of Israel intervened in the FX market.[11] We find (Table 6)

[10] The ex ante excess premiums are calculated as the difference between the stationary component of the nominal exchange rate (extracted using a Beveridge-Nelson decomposition) and the expected future nominal interest rate differentials. See Engel (2016).

[11] The advantage of using the dummy variable is its availability to the public.

that the baseline findings are essentially unchanged. The effect of purchases on the *annual* change of the exchange rate is insignificant. This implies that FX purchases did not directly affect the fundamentals of exchange rate dynamics. However, this does not rule out that FX interventions, as part of unconventional monetary policy, could indirectly affect the nominal exchange rate dynamics through their effect on inflation expectations.

(Table 6 about here)

4.4. Accounting for deviations from PPP and UIP

The results obtained for the expected relative PPP and UIP equations imply that for Israel, the interest rate differential (IRD) with the US is equal to the differences in expected inflation up to a constant and white noise. This implies ex-ante real interest rate parity. Therefore, it is useful to look at deviations from this constant and see whether these deviations result from deviations of the IRD from the equilibrium relationship or shocks to the differences in expected inflation (PPP). We assume that part of the constant difference between interest rates in Israel and the US is related to country risk measured by Israel's CDS premium. Part of it reflects constant lower liquidity in the Israeli bond market relative to the US. The remaining differences can be attributed to the real interest rate difference due to the differences in demographics and growth (Borio et al., 2017).

We estimated the following equation:

$$(11) \qquad (i_t - i_t^*) - (E\pi_{t+1} - E\pi_{t+1}^*) = (r_t - r_t^*) = \gamma + \delta CDS_t + \varepsilon_t$$

where γ captures the fundamental real interest rate differential and constant liquidity premiums.

Using the Bai-Perron multiple break test, we found that the equation is stable. This result strengthens the conclusions we draw from the separate estimates of the UIP and PPP relationships. It also suggests that the constant term is stable over the sample period. Next, we obtained the residuals and compared them with the IRD and PPP components. We tested for serial correlation using the Breusch-Godfrey Serial Correlation LM test and found a strong serial correlation. We can observe (Figure 5) that during some episodes, the residuals seem correlated mainly with the IRD. Indeed, the correlation of the residuals with the IRD is positive and significant. We then tested for the correlation of the residuals with the monetary policy by correlating them with the Bank of Israel and the Fed's policy rates. We found a stronger and more significant correlation of the residuals with the Bank of Israel than with the Fed policy rates.[12]

(Figure 5 about here)

These findings imply that some of the deviations were the product of monetary policy. A salient episode that shows up remarkably well in the data (Figure 5) is the Governor of the Bank of Israel's deal with the government to lower interest rates by two percentage points in December 2001. The move, which surprised markets, necessitated raising interest rates by more than five percentage points within six months. It is also not surprising that deviations from equilibrium characterized the crisis period—from October 2008 to the end of 2009—. Finally, it is interesting to note that toward the end of our sample—since the middle of 2017—we note a significant deviation. While it is too early to see how this episode will evolve, for now, the divergence of policy rates of the Bank of Israel and the Fed, while the difference in expected inflation is quite stable, may suggest that either the Bank of Israel is keeping interest rates too low or that the Fed is raising them too fast.

4.5. Testing the PPP and UIP using monthly rates of change

Engel (2016) confirmed the consensus in the literature that the UIP relationship does not hold for the short run. He tested the relationships using monthly rates of change in the exchange rate. We are in total agreement that testing UIP with long-term bonds as, for example, in Chinn and Meredith (2004), involves holding period risk. However, in this paper, we emphasize that the availability of financial assets that allows trading future inflation is crucial for the no-arbitrage assumption to hold. The shortest liquid maturity asset we have is twelve months' inflation breakeven rates. It is not even clear that if such assets existed, we could define a one-month change in the price level as inflation. Nevertheless, we follow Engel (2016) and take the 12th root of our annual inflation expectations series and interest rates and re-estimate equations (1) and (2).

We first estimate equations (1) and (2) for the period 1996 to 2007 and find (Table A3) that we cannot reject the hypothesis that PPP and UIP hold, even though the significance of our results is weakened relative to those obtained using annual data. We then estimate the equations for the entire sample, controlling for FX interventions, and find that our results still hold.

[12] Correlation coefficient of the residual with the Bank of Israel rate = 0.348.
Correlation coefficient of the residual with the Fed rate = 0.129.

4.6. Bootstrap Estimation

One advantage of Israel's data is that owing to its inflation experience, we have relatively, for advanced economies, a large variation in the regressors of the UIP and PPP regressions. This allows us to estimate the coefficients of interests more precisely. However,
it may be argued that our results may be driven by our sample choice that includes periods with trending disinflation (Figure 2). To rule out these selection issues, and in the way of robustness, we used bootstrap estimations. We estimated the Fama equation (equation 2) using bootstrap technique twice: for the whole sample and a subsample of 150 observations (out of 288). We used 100K replications and found similar results to the baseline estimation (Figure 6): the distribution of the Fama equation's coefficient is centered around 1 for both samples.

(Figure 6 about here)

Conclusions

We have shown that for Israel, a small open economy, we cannot reject the PPP and UIP conditions for the short run. Our results contrast with those arrived at in the literature. The difference between our findings and those of the literature may be driven by the characteristics of the Israeli economy and/or the data we used. Our results indicate that the existence of a market for inflation expectations for a short duration, namely 12 months, may help anchor the forward-looking PPP relationship. With free capital flows, the joint determination of the exchange rate on the current account (PPP) and the capital account (UIP) may help anchor the UIP relationship as well. Moreover, we believe that Israel is an advanced small open economy and is essentially a price-taker, contributes to the lower frictions in the trade of goods and services.

The finding that we cannot reject PPP and UIP for Israel suggests that monetary policy can be quite effective. The frictions pointed out by Rey (2016) and more recently by Plantin and Shin (2018), and Hofmann et al. (2020a, 2020b) are likely to play a smaller role: the expenditure switching effects of monetary policy probably dominate the balance sheet effects (a la Mundell-Fleming). We also show the absence of significant balance sheet effects of deviations from UIP or PPP. The effectiveness of monetary policy also means that monetary policy will continue to affect inflation expectations. Therefore, policymakers should bear in mind that policy mistakes, through expected PPP, may destabilize the dynamics of the exchange rate and may lead to financial instability.

Can we generalize from the case of Israel? To the extent that globalization proceeds and expands to the services industry, we will see fewer frictions, especially for advanced small open economies (Cavallo et al., 2018). This will likely increase the likelihood of finding evidence for the PPP relationship. To the extent financial markets in small open economies evolve and offer short term inflation swaps, the less likely it will be that arbitrage opportunities will develop in the foreign exchange market because the expected PPP relationship will anchor the expected exchange rate.

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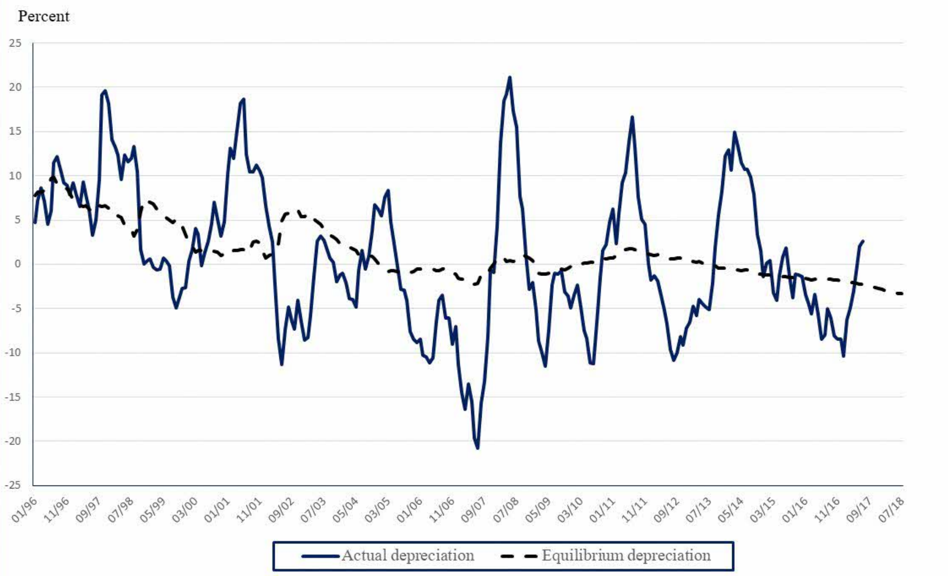
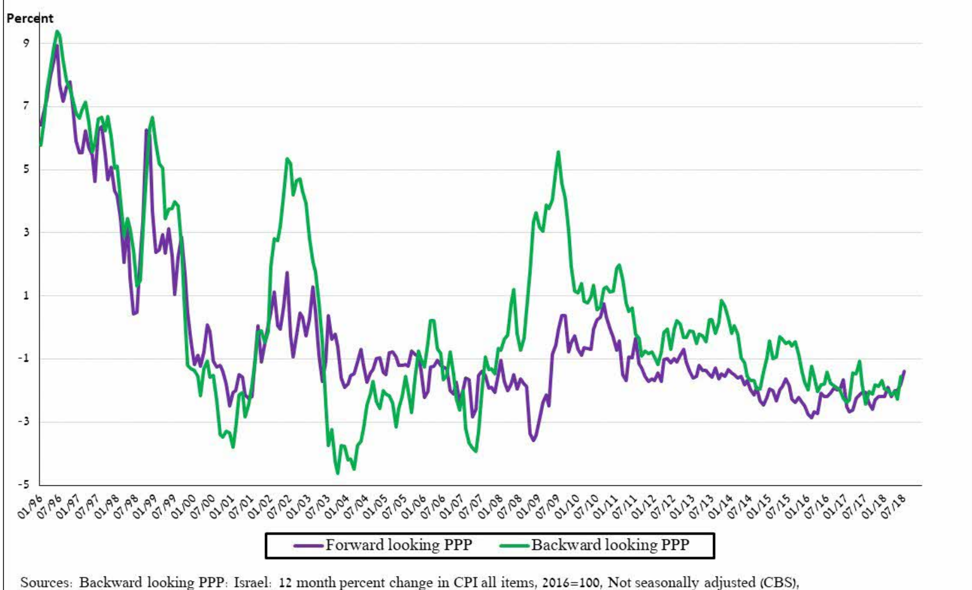
Figures and Tables

Figure 1
Actual versus Forecast Equilibrium Depreciation
Israel: 1996–2018



Notes: Actual deprecation: 12 months ahead change in N.I.S to $U.S exchange rate.
Equlibrium exchange rate: based on estimating a 2sls regression of the 12 months ahead change in the exchange rate on the yield differential between Israel and U.S 12 month bonds, instrumented by 12 months ahead difference in Israel and U.S expeceted inflation rate. See section 3.2.

Figure 2
Forward and Backward Looking Expected Inflation Differentials between Israel and the US.
(1996–2018)



Sources: Backward looking PPP: Israel: 12 month percent change in CPI all items, 2016=100, Not seasonally adjusted (CBS), U.S: CPI for All Urban Consumers: All Items, Index 1982-1984=100,, Not Seasonally Adjusted (FRED) Forward looking PPP: Israel: 12 months breakeven inflation rate, Bank of Israel. U.S: Michigan consumer survey, FRED.

Figure 3
The Yield on 1-year Bank of Israel Bonds and 1-year US Treasury bills (1996–2018)

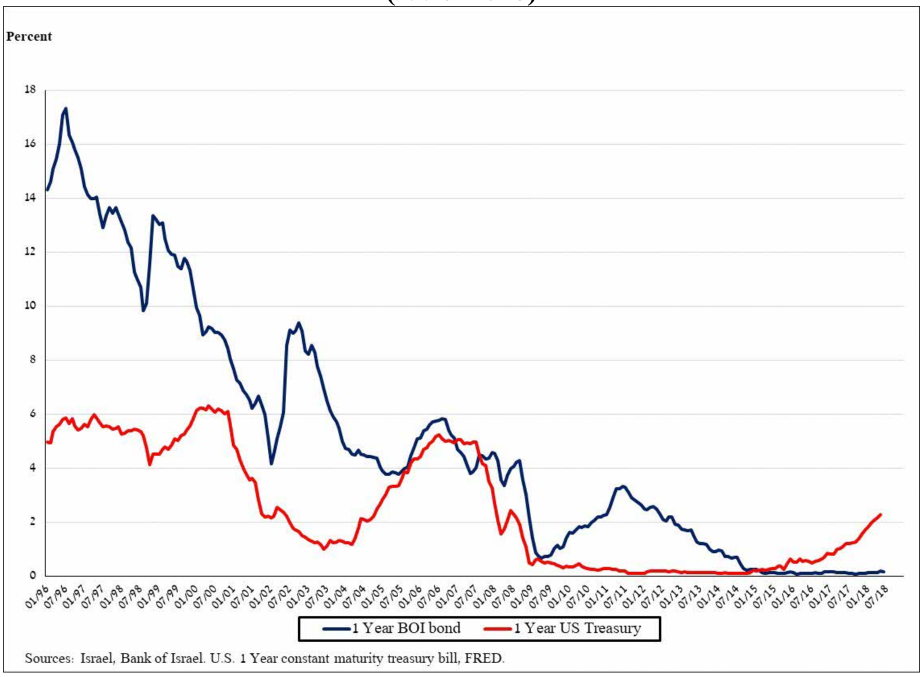


Figure 4
Five-Year NIS weekly bonds return (2008:1–2020:7)

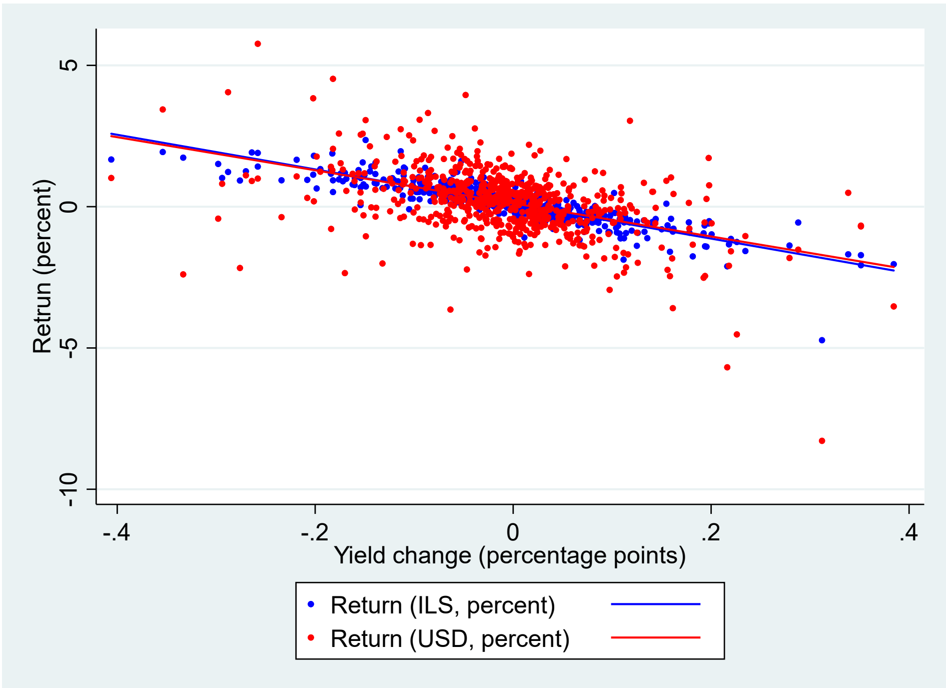
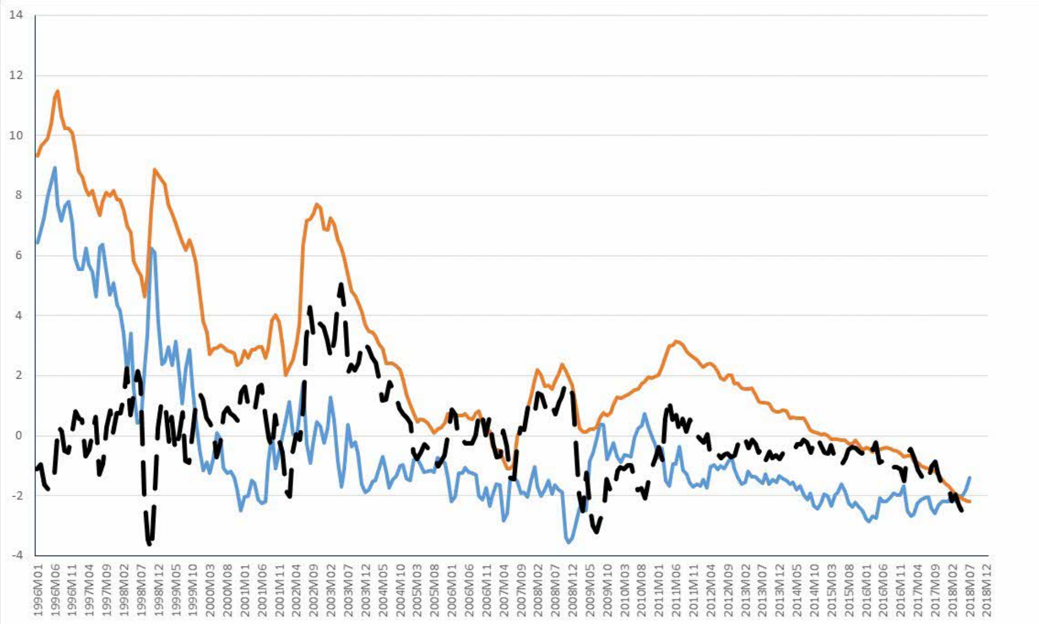


Figure 5

Deviations from Real Interest Rate Parity, the IRD, and PPP
(1996:1–2018:5)



Note: Residuals from: $(i_t - i_t^*) - (E\pi_{t+1} - E\pi_{t+1}^*) = \gamma + \delta CDS_t + \varepsilon_t$

Correlations: $r((i_t - i_t^*), \varepsilon_t) = 0.38$, $r((E\pi_{t+1} - E\pi_{t+1}^*), \varepsilon_t) = -0.09$

PPP IRD residuals

Figure 6
Bootstrap estimation results for Fama regression

Panel A: full sample (N=288) **Panel B: partial sample (N=150)**

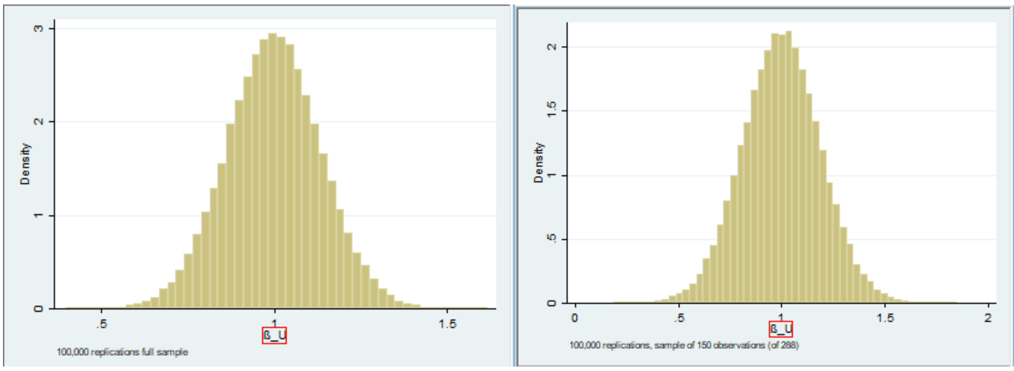


Table 1 – PPP equation: $(Es_{t+12} - s_t) = \alpha_P + \beta_P(E\pi_{t+12} - E\pi^*_{t+12}) + u_{P,t+12}$
1996:1–2017:7

α_P	95% interval	β_P	95% interval	$u_{P,t-1}$
1.060	(-2.40 , 4.540)	0.995	(0.277, 1.714)	-0.106 (0.022)
2SLS estimation				
1.064	(-0.882 , 3.001)	0.971	(0.359, 1.568)	

Notes: OLS: 95% confidence interval based o HAC standard errors & covariance (Pre-whitening with lags = 2 from SIC maxlags = 6, Bartlett kernel, Newey-West fixed bandwidth = 5)
VECM: no intercept or trend in CE or Var. 4 lags.
2SLS: Instruments $(\pi_{t-1} - \pi^*_{t-1})$, Fed policy rate lagged. Cragg-Donald F-stat: 382.

Table 2 – Fama regressions: $(Es_{t+12} - s_t) = \alpha_U + \beta_U(i_t - i^*_t) + u_{U,t+12}$
1996:1–2017:7

α_U	95% interval	β_U	95% interval	$u_{U,t-1}$
-2.032	(-6.593 , 2.528)	0.998	(0.162, 1.835)	-0.148 (0.028)
2SLS estimation				
-2.129	(-4.714 , 0.455	1.036	(0.473, 1.600)	

Notes: OLS: 95% confidence interval based o HAC standard errors & covariance (Pre-whitening with lags = 2 from SIC maxlags = 6, Bartlett kernel, Newey-West fixed bandwidth = 5)
VECM: no intercept or trend in CE or Var. 10 lags.
2SLS: Instruments, CDS, Bank of Israel rate lagged. Cragg-Donald F-stat: 711.

Table 3 – Capital account regression: $K_{i,t} = \alpha_{i,K} + \beta_{i,K} u_{U,t-j} + u_{K,t}$
1996:Q2–2017: Q4

	α_K	95% interval	β_K	95% interval
Nonresidents	875.8	(316.8 , 1434)	-83.5	(-150 , -16.8)
Residents	-1683	(-2340 , -1027)	95.8	(11.8 ,179)
Net flows	-771	(-1606 , 63.5)	26.8	(-83.7 ,137)

Notes: Capital account: portfolio + other investments.
OLS: 95% confidence interval based o HAC standard errors & covariance (Pre-whitening with lags = 0 from SIC maxlags = 4, Bartlett kernel, Newey-West fixed bandwidth = 4)
J=1 for nonresidents, j=3 otherwise.

Table 4 – Testing for the PPP puzzle
1996:1–2018:7

q		*s*		*log(p)-log(p*)*	
AR(1)	95% interval	AR(1)	95% interval	AR(1)	95% interval
0.980	(0.957 , 1.004)	0.978	(0.957 , 0.998)	0.974	(0.943 , 1.006)
Rolling AR(1) regression 24 months					
mean	Std. dev.	mean	Std. dev	mean	Std. dev
0.86	0.133	0.89	0.115	0.88	0.118
Rolling AR(1) regression 72 months					
0.953	0.040	0.949	0.032	0.958	0.041

q, *log(p)-log(p*):* OLS: 95% confidence interval based o HAC standard errors & covariance (Pre-whitening with lags = 1 from SIC maxlags = 6, Bartlett kernel, Newey-West fixed bandwidth = 5)
s: OLS: 95% confidence interval based o HAC standard errors & covariance (Pre-whitening with lags = 2 from SIC maxlags = 6, Bartlett kernel, Newey-West fixed bandwidth = 5)
Test for equal means 24 months Anova F-test (p=0.067), Welch F-test (p=0.077)
Test for equal means 72 months Anova F-test (p=0.037), Welch F-test (p=0.0037)

Table 5 – Testing for the UIP puzzle
1966:1–2017:7

$cov(E_t\rho_{t+1}, r_t^* - r_t) = -0.002; \; prob\ t = 0: \; 0.822$			
Modified Fama regression (Engel, 2016): $\rho_{t+1} = \zeta_s + \beta_s(i_t^* - i_t) + u_{s,t+1}$ 1996:1=2017:7			
ζ_s	95% interval	β_s	95% interval
0.024	(-0.002 , 0.050)	-0.000	(-0.005 , 0.005)
Real exchange rate regression (Engle, 2016): $q_t = \zeta_Q + \beta_Q(r_t^* - r_t) + u_{Q,t+1}$			
ζ_Q	95% interval	β_Q	95% interval
-2.233	(-2.262 , -2.204)	0.006	(-0.0007 , 0.014)

OLS: 95% confidence interval based on HAC standard errors & covariance (Bartlett kernel, Newey-West fixed bandwidth = 5)

Table 6 – Accounting for Bank of Israel FX interventions
1996:2–2017:7

PPP: $Es_{t+12} - s_t = \alpha_P + \beta_P(E\pi_{t+12} - E\pi^*_{t+12}) + \vartheta FX_t + \kappa FX_t^2 + u_{P,t+12}$					
α_P	95% interval	β_P	95% interval	FX_t	FX_t^2
0.642	(-3.494 , 4.779)	1.026	(-0.026 , 2.079)	3.303 (0.579)	-1.250 (-0.715)
$Es_{t+12} - s_t = \alpha_P + \beta_P(E\pi_{t+12} - E\pi^*_{t+12}) + \lambda D\_FX_t + u_{P,t+12}$					
α_P	95% interval	α_P	95% interval	D_FX_t	
0.249	(-4.163 , 4.660)	1.059	(-0.048 , 2.166)	2.547 (0.636)	
Fama regression: $(Es_{t+12} - s_t) = \alpha_U + \beta_U(i_t - i_t^*) + \vartheta FX_t + \kappa FX_t^2 + u_{U,t+12}$					
α_U	95% interval	β_U	95% interval	FX_t	FX_t^2
-3.769	(-7.670 , 0.161)	1.298	(0.386 ,2.210)	5.495 (1.345)	-1.727 (-1.242)
Fama regression: $(Es_{t+12} - s_t) = \alpha_U + \beta_U(i_t - i_t^*) + \lambda D\_FX_t + u_{U,t+12}$					
α_U	95% interval	β_U	95% interval	D_FX_t	
-5.377	(-10.682 , -0.072)	1.565	(0.670 ,2.459)	5.330 (1.595)	

Notes: 2SLS: 95% confidence interval based on HAC standard errors & covariance (pre-whitening with lags from AIC max lags, (Bartlett kernel, Newey-West fixed bandwidth = 5)
FX: monthly billions of $US purchased. D_FX: dummy for months with FX intervention.
t statistics for FX and D_FX in parenthesis,
2SLS: Instruments PPP in parentheses,
2SLS: Instruments PPP $\pi_{t-1} - \pi^*_{t-1}$, Fed policy rate lagged.
2SLS: Instruments UIP: CDS, Bank of Israel rate lagged.

Appendix

Tables

Table A1 – PPP equation alternative specifications:
1996:1-2017:7

OLS: $(Es_t - s_{t-1}) = \alpha_P + \beta_P(\pi_{t-1} - \pi^*_{t-1}) + u_{P,t}$				
α_P	95% interval	β_P	t-statistic	95% interval
0.520	(-3.752 , 4.791)	0.553	0.848	(-0.730 , 1.838)
2SLS: $s_t - s_{t-12} = \alpha_P + \beta_P(\pi_t - \pi^*_t) + u_{P,t+1}$				
0.007	-(0.027 , 0.041)	1.194	1.833	(-0.089 , 2.47)

Notes: OLS: 95% confidence interval based on HAC standard errors & covariance (Pre-whitening with lags=3 from SIC, max lags=6, Bartlett kernel, Newey-West fixed bandwidth = 5)
2SLS: 95% confidence interval based on HAC standard errors & covariance (Pre-whitening with lags from SIC, Bartlett kernel, Newey-West fixed bandwidth = 5)
2SLS: Instruments $\pi_{t-12} - \pi^*_{t-12}$, Fed policy rate (t-13) Bank of Israel rate (t-13). Cragg-Donald F-stat: 32. 8
2SLS estimated for 1977:2–2018:6.

Table A2 – Granger causality tests of deviations from PPP equations on the capital account of the balance of payments:
1996:1-2017:7

OLS: $(Es_{t+12} - s_t) = \alpha_P + \beta_P(\pi_{t+12} - \pi^*_{t+12}) + u_{P,t+12}$				
Granger causality tests of $u_{P,t+12}$ on $K_{i,t}$				
	Test	N	F-statistic	Prob.
Nonresidents	$K_{i,t}$ does not Granger cause $u_{P,t+12}$	258	2.010	0.157
	$u_{P,t+12}$ does not Granger cause $K_{i,t}$	258	5.313	0.022
Residents	$K_{i,t}$ does not Granger cause $u_{P,t+12}$	257	0.911	0.4033
	$u_{P,t+12}$ does not Granger cause $K_{i,t}$	257	4.053	0.018

Notes: 1 and 2 lags used in tests, respectively.

Table A3 – PPP and Fama regression monthly rate of change

<table>
<tr><td colspan="6">PPP: $(Es_{t+1} - s_t) = \alpha_P + \beta_P(E\pi_{t+1} - E\pi^*_{t+1}) + u_{P,t+1}$
OLS: 1996:2-2007:12</td></tr>
<tr><td>α_P</td><td>95% interval</td><td>β_P</td><td>95% interval</td><td></td><td></td></tr>
<tr><td>-0.001</td><td>(-0.003 , 0.004)</td><td>1.029</td><td>(0.129 , 1.93)</td><td></td><td></td></tr>
<tr><td colspan="6">PPP: $(Es_{t+1} - s_t) = \alpha_P + \beta_P(E\pi_{t+1} - E\pi^*_{t+1}) + \vartheta FX_t + \kappa FX_t^2 + u_{P,t}$
OLS: 1996:2-2017:7</td></tr>
<tr><td>α_P</td><td>95% interval</td><td>α_P</td><td>95% interval</td><td>FX_t</td><td>FX_t^2</td></tr>
<tr><td>-0.001</td><td>(-0.003 , 0.003)</td><td>0.952</td><td>(-0.060 , 1.965)</td><td>0.006
(1.247)</td><td>-0.002
(-1.179)</td></tr>
<tr><td colspan="6">Fama regression: $(Es_{t+1} - s_t) = \alpha_U + \beta_U(i_t - i_t^*) + u_{U,t+1}$
2SLS: 1996:2-2007:12</td></tr>
<tr><td>α_U</td><td>95% interval</td><td>β_U</td><td>95% interval</td><td></td><td></td></tr>
<tr><td>-0.003</td><td>(-0.009 , 0.003)</td><td>1.268</td><td>(-0.079 , 2.615)</td><td></td><td></td></tr>
<tr><td colspan="6">Fama regression: $(Es_{t+1} - s_t) = \alpha_U + \beta_U(i_t - i_t^*) + \vartheta FX_t + \kappa FX_t^2 + u_{U,t+1}$
2SLS: 1996:2-2007:12</td></tr>
<tr><td>α_U</td><td>95% interval</td><td>β_U</td><td>95% interval</td><td>FX_t</td><td>FX_t^2</td></tr>
<tr><td>-0.005</td><td>(-0.011 , -0.000)</td><td>1.829</td><td>(0.459 ,3.199)</td><td>0.009
(1.834)</td><td>-0.003
(-1.803)</td></tr>
<tr><td colspan="6">Notes: 2SLS: 95% confidence interval based on HAC standard errors & covariance (pre-whitening with lags from SIC max lags, (Bartlett kernel, Newey-West fixed bandwidth = 5)
FX: monthly billions of $US. purchased. D_FX: dummy for months with FX intervention.
t statistics in parenthesis for FX and D_FX,
2SLS: Instruments UIP, CDS, Bank of Israel rate lagged.</td></tr>
</table>

Appendix 1: Estimation of Israel's credit risk before 2002

Data on Israel's credit default swap (CDS) is available from Bloomberg only from July 2002. Since Israel was considered an emerging market, we regressed Israel's CDS on the EMBI index. We used the estimated equation to generate estimates for Israel's CDS before July 2002. The R squared obtained was 0.8. Figure A1below shows the actual CDS and the estimated CDS.

Figure A1
Actual and estimated 5 CDS for Israel: 2002:07-2007:12

Sources: Israel 5 year CDS and EMBI index: Bloomberg
Fitted values from the regression $CDS_t = \rho + \tau EMBI_t + \upsilon_t$. Estimated using OLS with HAC standard errors & covariance (Pre-whitening with lags = 1 from SIC, max lags = 3, Bartlett kernel, Newey-West fixed bandwidth = 4). R^2=0.81.

Appendix 2: Unit root tests for main variables

In this appendix, we report the results of the unit-root test on the main variables used in our analysis: the rate of deprecation of the currency, the 12 months' bond spread, the difference in expected inflation, Israel's risk premium, and FX purchases as well as the standard deviation of the exchange rate. We find that we can reject both that series have unit root and the hypothesis of a joint unit root process.

Table A4 – Unit root tests

Group unit root test: Summary

Series: Exchange rate depreciation, Yield spread, Expected inflation differential, 5Y CDS, FX purchases, Std error.

Sample: 1995M 01 2018M 12

Exogenous variables: Individual effects

Automatic selection of maximum lags

Automatic lag length selection based on SIC: 1 to 13

Newey-West automatic bandwidth selection and Bartlett kernel

Method	Statistic	Prob.**	Cross-sections	Obs
Null: Unit root (assumes common unit root process)				
Levin, Lin & Chu t*	-1.94800	0.0257	6	1570
Null: Unit root (assumes individual unit root process)				
Im, Pesaran and Shin W-stat	-4.93007	0.0000	6	1570
ADF - Fisher Chi-square	53.1993	0.0000	6	1570
PP - Fisher Chi-square	100.260	0.0000	6	1594

** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality