Essays in Financial Economics: Terror, Consumption, and Investment, Currency Options and Liquidity Premium, and Purchasing Power Parity

by

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Submitted to the Department of Economics in Partial Fulfillment of the Requirement for the Degree of

Doctor of Philosophy

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Abstract

This thesis is composed of three chapters, each includes one paper. The first chapter includes a paper that analyses the impact of terror on consumption and investment. This paper provides evidence on how consumers and investors react to terror attacks based on a new database from the Israeli-Palestinian conflict. An increase in terror casualties triggers households to alter their perceived personal security and expected future income. Only ex-post do households distinguish a temporary from a permanent increase in terror casualties. A temporary increase in the number of terror casualties causes a bust-boom cycle of durables consumption and irreversible investment; nondurables are affected less. A permanent increase in the number of terror casualties causes a one-time drop in consumption. This is in line with the theory on irreversible investment and durables consumption: terror generates temporary uncertainty about personal security and future income, which in turn causes a bust-boom cycle of durables due to bunching of purchases in later periods. A permanent increase in terror causes neither bunching nor boom. Similar results are obtained for the effect of terror casualties on fixed capital.

The second chapter includes a paper titled: "Arbitrage Tests of Israel's Currency Options Markets." The aims of this study are threefold. First, we test the validity of the Black and Scholes (B-S) model as a naive option-pricing model for the case of an exchange-rate target zone. We find that although we cannot reject the weakly efficient market hypothesis (except for very-near-maturity deep-ITM options), we can reject the strongly efficient market and/or the B-S model validity hypotheses. The banking sector could have utilized arbitrage opportunities, notably for out-of-the-money, at-the-money, and far-from-maturity options, especially when employing inter-temporal weighted-average implied standard deviation. Second, we estimate the liquidity premium for currency options by using a unique data set that allows us to comparing tradable and non-tradable options. The liquidity premium, though positive in average, is found to be negative for some options. This is an indication that there could have been arbitrage opportunities, especially for the banking sector. Third, we examine the null hypothesis that the Israeli currency options market is efficient, an issue that has not been investigated. Ex-post tests of arbitrage and dominance conditions do not permit rejection of the null hypothesis, except for very-near maturity, deep-in-the-money (ITM) options. The paper enhances the literature by using a unique database from the Israeli currency options market, which includes currency options traded on the Tel Aviv Stock Exchange and (non-tradable) Bank of Israel currency options. In addition, this paper examines B-S when the exchange rate is confined to a target zone.

The third chapter includes a paper that analyses the robustness of exchange rate models, unit roots and cointegration. Three basic models have been proposed to explain the exchange rate: Purchasing Power Parity (PPP), the Balassa-Samuelson model and the random walk model. The robustness of these models is not merely a statistical curiosity but has important implications in many economic and financial models. During the last two decades this issue has attracted many researchers who have used the latest econometric methods. However, there are many who are skeptical and the robustness of the PPP theory is still an open question. The purpose of this paper is threefold: first, to summarize the empirical literature on modeling the exchange rate process; second, to suggest an appropriate transformation of the aforementioned models into an econometric specification; and finally to introduce a precise framework for examining their separate or "pooling" validity. In particular, I use a formal Complete Vector Error Correction Model (CVECM) with several convincing statistical methods. Applying this methodology to the 1920s float, I find evidence that cannot reject the hypothesis of the invalidity of the long run PPP with short run deviations.

JEL classification: E21; F31; G13

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Acknowledgement

There are many people whom I would like to thank for helping me throughout my stay at MIT. I want to thank all of my professors, especially my three advisors: Olivier Blanchard, Whitney Newey, and Jerry Hausman. They all were extremely supportive and showed a great interest in my work. Their comments and suggestions greatly improved this dissertation. Their contagious good mood and generosity made my stay at MIT both rewarding and fun. I am grateful for their generosity in providing financial support. Olivier's comments and corrections, especially to my mangled grammar, on the drafts have been much appreciated. Whitney's high standards led to many improvements in this thesis and always challenged me to be as rigorous as possible. Jerry has been a tremendous source of encouragement and common sense. I want to thank Simon Benninga for his special advice in regard to my second paper of this dissertation. I am also thankful to Guido Kuersteiner and Joseph Zeira and two anonymous reviewers for their helpful advice and comments. I would also like to thank the participants in the macroeconomics, econometrics, and finance workshops and seminars at MIT for useful observations. I would like to thank Simcha Bar-Eliezer from Israel CBS, Jakob Bruade from the Bank of Israel, Tomer Gardi from Btselem, Soli Peleg from Israel CBS, and Yael Shahar from ICT for providing data and helpful information. Special acknowledgement with regard to the second paper is included in the first footnote of Chapter 2.

I am indebted to Joshua Angrist, Simon Benninga, Dani Tssidon, and Joseph Zeira, who stimulated me to continue studying economics at MIT and supported me through the end of my PhD. I am thankful to my friends at Herzog, Fox, and Neeman, Advocates, and Deloitte Touche in Tel-Aviv for encouraging me to complete my thesis.

I have special thanks to the US State Department and the US Embassy in Tel-Aviv for their generous scholarship and the MIT Department of Economics for several grants.

Last but not least, I am grateful to my wife Adan who dedicated her time solely for me. She supported and encouraged me throughout the writing of my dissertation. Adan kept the enthusiasm going anytime the future looked uncertain and the present felt painful.

For *Adan* and *Saji*

إلى أحباء قلبي زوجتي عدن وإبني ساجي

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- Supervised over 1,000 students and monitored their academic progress
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1996 **Ben-Porath Scholarship** (honor researcher)

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1993 **Gelfat Award** (excellent graduate student)

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SELECTED PUBLICATIONS AND PRESENTATIONS

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Table of Contents

Chapter 1: Terrorizing consumers and investors

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Abstract

This paper provides evidence on how consumers and investors react to terror attacks based on a new database from the Israeli-Palestinian conflict. An increase in terror casualties triggers households to alter their perceived personal security and expected future income. Only ex-post do households distinguish a temporary from a permanent increase in terror casualties. A temporary increase in the number of terror casualties causes a bust-boom cycle of durables consumption and irreversible investment; nondurables are affected less. A permanent increase in the number of terror casualties causes a one-time drop in consumption. This is in line with the theory on irreversible investment and durables consumption: terror generates temporary uncertainty about personal security and future income, which in turn causes a bust-boom cycle of durables due to bunching of purchases in later periods. A permanent increase in terror causes neither bunching nor boom. Similar results are obtained for the effect of terror casualties on fixed capital.

Keywords: Consumption; Investment; Uncertainty; Terror

JEL classification: E21; E27; G31; Z00

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1. Introduction[*, 1]

While terror is not new, the effect of terror on consumption and investment has not been empirically examined in prior economic studies. A major effect of terror is the creation of uncertainty *vis-à-vis* personal security as well as future income.[2] Against this background, this paper addresses the following questions:

(1) What are the impulse response functions of consumption and investment?
(2) Are the impact response functions for durables, nondurables, and services different?
(3) Is there a long-term effect of terror on consumption and investment?
(4) Do consumers develop a hedonic adaptation to terror?

Answers will assist policymakers to understand how terror affects consumption and investment, which together constitute significant portions of GDP, and perhaps to neutralize the effects of terror in a timely fashion to prevent destabilization of the economy.

The above questions are addressed using a new database on Israeli consumption and terror casualties from the Israeli-Palestinian conflict for the period 1980-2002. As Figure 1 shows, the Israeli death toll fluctuated in terms of severity and frequency[3]. Terror was likely a leading cause of the Israeli economy's recent worst economic performance, although these years also coincided with end of the high-tech bubble that also adversely affected the Israeli economy. In 2001 and

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[*] I am grateful to Olivier Blanchard, John Hassler, Jerry Hausman, Guido Kuersteiner, Whitney Newey, and Joseph Zeira for their helpful advice and comments. I would also like to thank the participants in macroeconomics, econometrics, and finance workshops and seminars at MIT for useful observations. I would like to thank Simcha Bar-Eliezer from Israel CBS, Tomer Gardi from Btselem, Soli Peleg from Israel CBS, and Yael Shahar from ICT for providing data and helpful information. All remaining errors are mine.

[1] This paper may include political terms. These are given only for the purpose of economic analysis and there is absolutely no political presumption in their usage.

[2] Systematic risk will be taken to be synonymous with uncertainty regarding future events of relevance to consumers and investors.

[3] See section 3.2 for further details on the pattern of terror casualties in terms of number of dead and wounded and type of target.

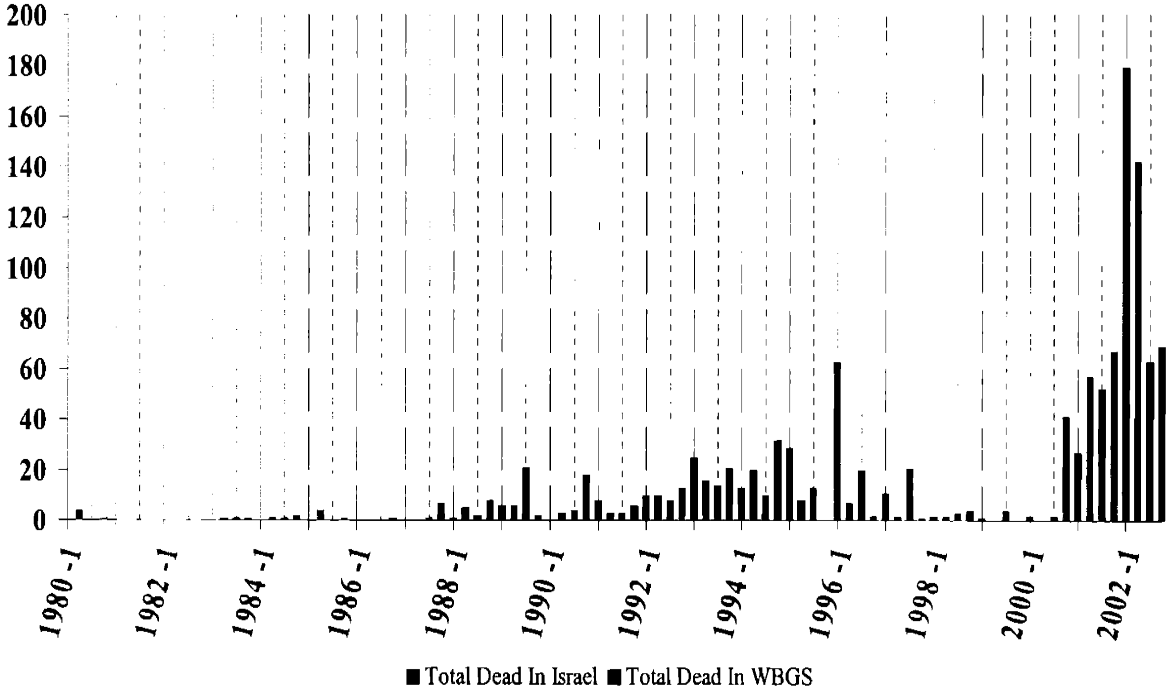

Figure 1: Total Israeli dead in terror attacks against Israelis in Israel and West Bank and Gaza Strip

Source: Database was compiled by the author from various sources listed in section 3.1.

2002 GDP decreased two years in a row for the first time since 1953, private consumption declined in 2002 for the first time since 1980, and the annual inflation for 2003 was negative for the first time since 1948.

Section 2 introduces the theoretical background. Section 3 analyzes the pattern of terror in the Palestinian- Israeli conflict. I identify two types of terror attacks with different implications for perceived uncertainty with respect to personal security and future income. I then analyze the different effects of these uncertainties on consumption and investment. The analysis focuses on aggregated and disaggregated consumption data-- durables, nondurables, and services as well as fixed capital and inventories. Section 4 concludes.

2. Theoretical background

Macroeconomic effects of terror have been studied by Abadie and Gardeazabal (2003), Eldor and Melinck (2004), Enders and Sandler (1996, 1991), Fielding (2003), and Fleischer and Buccola (2002). Eldor and Melinck (2004), Fielding (2003), and Fleischer and Buccola (2002) in particular also focus on terror against the population in Israel. These empirical studies have investigated the reduced form effect of terror on economic variables such as GDP, foreign investment, saving, financial markets, and tourism. In contrast, this paper considers the effect of terror on aggregate consumption and investment.[4]

Terror is expected to affect households' decisions on consumption and firms' decisions on investment through consequences for personal security and future GDP. The effect of terror on personal security is especially important when terror hits civilians in the course of their everyday lives in a random pattern, as has occurred in Israel. Terror affects GDP through two main mechanisms: increased uncertainty over future GDP growth (higher variance) and depressed expected GDP growth (lower expected mean).

2.1. Terror increases uncertainty

Terrorists intimidate people by threatening their wealth and lives (no matter what is their motive and whether legitimate or not). They do so by causing casualties through attacks against civilian and military targets. In many cases, the population cannot ex-ante learn whether a change in the distribution of casualties has taken place; people only ex-post infer the change in uncertainty by observing terrorists acts, government responses, and outcomes including terror casualties. An increase in the number of casualties increases uncertainty that the population faces.[5] This increase takes place through three channels:

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[4] After finishing writing this paper, I learned of Eckstein and Tsiddon (2004), which is a working paper that studies the impact of terror on the Israeli consumption.

[5] In some case, a sudden change of attack method is sufficient to bring about a new assessment of the uncertainty level, even if there was no change in terror casualties, e.g., the use of anthrax in the U.S. in late 2001. As will be clear in Section 3 below, in the case of Israel, the number of casualties was a sufficient indicator for a possible change in uncertainty.

1. Increasing contemporaneous uncertainty: There is a positive correlation between the number of terror casualties and the uncertainty that people face during the contemporaneous quarter.

2. Shifting the probability of future terror casualties upward (greater expected mean of terror casualties): The people' posteriors are usually centered around the contemporaneous level of terror casualties.

3. Making the future less certain (increasing the variance of the distribution of future terror casualties): There is a positive correlation between the mean of current casualties and the variance of future casualties. There are two reasons for this observation. First, terror casualties are bounded from below by zero and unlimited from above. Second, the higher the current number of terror casualties is, the faster the terror and counter-terror may escalate. Terror may escalate due to temptation and higher morale on part of the terrorists, whereas counter-terror may escalate due to public support and lower tolerance for terror attacks on part of the government. Such escalations may increase, offset, or indeed decrease the terror threat in the short run.

If terror casualties persist at the new high level without further escalation on part of the government or the terrorists, the state of terror will appear stable; hence the perceived uncertainty decreases.

This uncertainty with respect to actual evolution of terror casualties, perhaps most importantly, generates in turn uncertainty over future personal security and GDP growth through the following three channels:

1. Economic instability: As shown in Abadie and Gardeazabal (2003), GDP is sensitive to the level of terror casualties. Since the path of terror casualties is uncertain, the path of the aggregate supply and demand is uncertain as well. Also, the impact on GDP is not immediate and lasts for some period over which the magnitude and timing of the impact is uncertain.

2. Political instability: A failure to counter terror or an unpopular move by the government will decrease its credibility and shock its stability.[6] GDP growth rates, as is well documented, are negatively impacted by measures of political instability.[7]

3. Personal insecurity: That terror decreases personal security is self-evident. This increase in insecurity prompts the consumers and investors to undertake measures to protect themselves, such as decreasing their use of public transportation, shopping in malls, eating out in restaurants, and perhaps attending weddings and other family celebrations.

I shall address the macroeconomic effects of terror by analyzing the *change in*, rather than the *level of*, the number of casualties. The former calls on consumers and investors to alter their perceived uncertainty.[8] The effect of uncertainty on durables consumption and irreversible investment is immediate and can be substantial, as will be evident from section 3.3.

2.2. Depressing GDP growth

On the one hand, an increase in the number of terror casualties will depress GDP growth; the longer this unanticipated shock is expected to last, the stronger the effect will be on the GDP growth rate. On the other hand, such an increase might lead the government to increase resources directed at preempting terror and reduce terror to a level lower than prior to the recent attack, in which case the growth rate of GDP will increase. The answer to which effect will dominate is, in general, unclear. However, whichever effect dominates, its impact will be immediate: according to rational expectations theory and the permanent income hypothesis, forward-looking agents immediately incorporate the anticipated change in GDP growth rate into their decisions. Such a

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[6] Terror attacks in Israel over the past 15 years, for example, caused a change of a prime minister every two years on average, and no previous prime minister was successful in a subsequent election. The wars on the regime of Saddam Hussein in Iraq and the subsequent terror were political issues in the U.S. and the U.K.

[7] See the literature on the effect of political conflict on economic variables, e.g., Acemoglu and Robinison (2001), Alesina et al. (1996), and Barro (1991).

[8] Abadie and Gardeazabal (2003), Enders and Sandler (1996, 1991), and Fleischer and Buccola (2002) considered only the effect of the level of casualties.

change is expected to be positively correlated with durables and nondurables consumption as well as reversible and irreversible investment.

While effects of a higher uncertainty and a lower GDP growth rate are expected to be immediate, the former affects mainly durables consumption and irreversible investment, while the latter affects all consumption and investment. This difference will prove helpful in distinguishing which effect is at play.

2.3. The impact of uncertainty

2.3.1. Temporary uncertainty

Leahy and Zeira (2003), in a general equilibrium model that explicitly examines the timing of durables and nondurables consumption, observed that a temporary shock to wealth or income causes a bust-boom cycle in durable spending due to bunching of purchases in later periods. As consumers accumulate wealth, demand for durables returns to its pre-shock level. In the case of permanent shocks, there is a one-time reduction in durables consumption; there is no bunching and no boom. They also observe that wealth shocks have ambiguous effects on nondurables consumption.[9]

An alternative approach to examining the effects of uncertainty on durables is through irreversible investment theory. In the short-run, durables consumption, just like irreversible investment, is characterized by durability, irreversibility, and indivisibility.[10] Therefore, its response in the short-run to aggregate shocks is more in line with the aggregate implications of lumpy investment models than with the predictions of a frictionless life cycle-permanent income hypothesis (LC-PIH) such as Hall (1978) and Flavin (1981), especially in periods with high aggregate uncertainty (see Bar-Ilan and Blinder, 1988 and 1992; Bernanke, 1984 and 1985; Lam,

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[9] When the price of durables relative to nondurables is assumed exogenous, changes in wealth cause a delay in the purchases of durables without changing their quantities, such changes have no effect on the quantities or timing of nondurables. This is what Leahy and Zeira called the perfect "insulation effect." In the general case, when this assumption is relaxed, the effect on nondurables is ambiguous.

[10] Durability separates consumption from purchase, irreversibility means long-term commitment, and indivisibility implies lumpy and discontinuous purchases.

1989; Mankiw, 1982).[11] Therefore, models that relate uncertainty to irreversible investment should be applicable in a straightforward manner to the relationship between uncertainty and consumption.

Bernanke (1983) and Cukierman (1980) have demonstrated the option value of deferring irreversible investment when there is temporary uncertainty and when information pertinent to the choice of investment arrives over time and makes the future less uncertain. They showed that a temporary increase in uncertainty could cause a bust-boom cycle in irreversible investment spending. In an application of this model to durables consumption, we should expect a similar pattern of response by consumers to temporary increases in uncertainty due to increases in terror victims.

2.3.2. Permanent uncertainty

There are two principal different, but not mutually exclusive, models relating a permanent increase in uncertainty with irreversible investment, which can be applied to examine the relation of a permanent increase in uncertainty with durables consumption: the value of waiting and the (S,s) models. A model that directly examines the effect of a permanent increase in uncertainty on consumption is the precautionary saving model, as explained below.

Pindyck (1991) developed a model related to Bernanke (1983) and Cukierman (1980), in which information arrives over time but the future is always uncertain. Here too there is a value associated with waiting and a permanent increase in uncertainty of the economic environment has a negative long-term effect on the amount of actual irreversible investment, which should be symmetrically applicable to durables consumption.

Bar-Ilan and Blinder (1988, 1992), Bertola and Caballero (1990), Caballero (1993), and Eberly (1994) applied the (S,s) model to durables purchases. They concluded that higher risk broadens

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[11] Note that Caballero (1990) validates the LC-PIH for the medium-long-run. Several justifications for the slow adjustment of durables consumption have been advanced in the literature, including inter-temporal complementarities in consumption, convex adjustment costs, and financial markets imperfection. Surveys of the modern consumption literature by Deaton (1992) and Hall (1989) discuss some of the efforts that have been made to explore departures from the frictionless LC-PIH.

the (S,s) bands, so durables deviate more from their optimal levels before being adjusted. Hence, a permanent increase in uncertainty causes a drop in fixed investment as well as durables consumption.[12]

The precautionary saving model directly examines the effect of permanent increases in uncertainty on consumption. Under inter-temporal optimization and decreasing CARA utility function, an increase in uncertainty leads consumers to defer consumption (to be more prudent and increase precautionary saving). Put differently, when risk increases, consumption declines, after which its growth rate increases without a bust-boom cycle in the short-run. This result is reinforced if it is also assumed that consumers have adaptive expectations, which is likely when consumers consider the state of terror.[13] If changes in subjective uncertainty fall more heavily on some individuals, as is likely the case in terror attacks (for example, children and workers using public transportation), their aggregate effect can be much larger (Blanchard and Mankiw, 1988). Therefore, if changes in terror casualties are interpreted as a permanent increase in uncertainty, we should expect consumption to decline in the short run before it adjusts to its long-term level.

Thus, a permanent increase in uncertainty causes a decline in irreversible investment and durables consumption in all three models: the value of waiting, the (S,s), and the precautionary saving models.[14] However, there is an important difference in implications between the value of waiting and the precautionary saving models: the former is relevant only for durables consumption and irreversible investments, whereas the latter is relevant to nondurables and inventories as well as durables consumption and fixed capital. This difference in reaction to

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[12] The (S,s) type of models generally considers only permanent shocks.

[13] In such a case, consumers expect further decline in their income, which enhances their precautionary saving.

[14] The depressing effect of uncertainty is well-documented in the theoretical literature. Ingersoll and Ross (1988) have examined irreversible investment decisions when the interest rate evolves stochastically, but future cash flows are known with certainty. As with uncertainty over future cash flows, the uncertainty regarding future interest rate creates a value to waiting. Investing is depressed further as the volatility of interest rates grows. Caballero and Corbo (1988) have shown how uncertainty over future real exchange rates can depress exports. Dornbusch (1987) has noted that uncertainty over future tariffs structure creates an opportunity cost to committing capital to new physical plants.

uncertainty will be useful when empirically distinguishing irreversibility effects from precautionary saving effects.

In all of the above models, the effect of macro-uncertainty, whether temporary or permanent, on aggregate irreversible spending follows from the micro-analysis. Terror attacks are a classic case where uncertainty about future macro-information pertinent to the choice of durables consumption and irreversible investment is not eliminated by aggregation level and the law of large numbers does not apply.[15]

3. Empirical findings

Due to the lack of data availability that relates macro-uncertainty with durables consumption and irreversible investments, little empirical evidence is provided in the literature for the effect of macro-uncertainty on durables consumption and irreversible investment. A notable exception is Romer (1990). She suggested, based on Bernanke (1983), that the fall in durables consumption in the year following the October 1929 Great Crash could be best explained by the temporary increase in uncertainty regarding the course of future income.[16] That the Great Crash caused uncertainty was evident from the decline in surety expressed by contemporary forecasters. Romer's case study included only one event and she did not detect a bust-boom cycle in consumption.[17] Here I examine the effect of changes in macro-uncertainty on durables consumption over 23 years, a period over which macro-uncertainty fluctuated due to changes in terror fatalities.

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[15] For examples on idiosyncratic shocks that eliminate cycles, see the (S,s) models in Caballero (1993), Caballero and Engle (1991), Caplin and Spulber (1987), and Zeira (1990).

[16] Romer concluded, based on informal argument, that Bernanke's (1983) model implies that uncertainty is positively correlated with nondurables consumption. This conclusion seems erroneous, as illustrated above. Further, Bernanke (1985) found that durables and nondurables are neither strong substitutes nor strong complements; thus the 'spillover' effect from slowly adjusting durables to nondurables may not be especially important in practice. In fact, Romer, too, found that uncertainty had had an ambiguous effect on nondurables consumption, a finding that is supportive of our conclusion above.

[17] Hassler (2001), following Romer (1990), relates automobile purchases to uncertainty that can be learned from fluctuations in the exchange.

3.1. Data description

For the purposes of this study, I constructed a new daily database that provides an accurate and complete coverage of politically-motivated terror attacks carried out by Palestinians against Israeli targets during the years 1980-2002.

This database is based on the following websites: ABC News; ADL; Al-Jazeera (Arabic newspaper); B'Tselem (The Israeli Information Center for Human Rights in the Occupied Territories); BBC News World Edition; CNN; GPO - Government Press Office; Haaretz (Israeli Hebrew newspaper); HAMAS (Islamic Resistance Movement); ICT - The International Policy Institute for Counter-Terror; IDF - Israel Defense Forces; Islam Online; Jerusalem Post (Israeli English Newspaper); LAW (The Palestinian Society for the Protection of Human Rights and the Environment); MFA - Ministry of Foreign Affairs; Peace Now; PIC - Palestinian Information Center; Shia News; The Ministry of Labor and Social Affairs and the National Insurance Institute; TVA - Terror Victims Association; Walk for Israel; and others.

The data includes 687 observations. Each observation includes the following information: date, town, number killed, number injured, method of attack, type of target, and the particular Palestinian organization carrying the attack. The details (and most importantly, the number of fatalities) of each observation in the data was confirmed by at least two sources with the exception of those observations recorded by B'Tselem. The reason for this exception is that B'TSelem, an independent human rights organization, ensures the reliability of information it publishes by conducting its own fieldwork and research, whose results are thoroughly cross-checked with relevant documents, official government sources, and information from other sources, among them Israeli, Palestinian, and other human rights organizations.

For all observations, except five, the number of reported fatalities was identical for all sources. For those five observations, I use the number of fatalities that was recorded by more sources (the discrepancy was less than 2 fatalities).

In some cases, the body of kidnapped soldier was found in a quarter after the quarter during which the kidnapping happened, or some of the casualties died later of their wounds and their day of death fell in a quarter following the quarter during which the attack occurred. In such cases, the day of the kidnapping or attack was used.

I maintained a distinction between Israel within its 1948 borders and West Bank and Gaza Strip (WBGS) that was occupied by Israel in 1967. I also distinguish between Jewish and Arab neighborhoods of Jerusalem. A terror attack on Jews that took place in East Jerusalem was classified as "in Israel" if the attack occurred in a Jewish neighborhood, and "in the West Bank" if Jews were attacked in an Arab neighborhood.

Data on national accounts were provided by the Israeli Central Bureau of Statistics.[18] All national accounts data employed were chained at 2000 prices.

3.2. The pattern of terror

Terror attacks against the population of Israel have had ebbs and flows in terms of severity and frequency (see Figure 1). Between January 1980 and December 2002, terror left a total of 914 dead and 4,755 wounded in attacks against civilians, in addition to 294 dead and 515 wounded in attacks against military personnel (see further details in Table 1). Terror attacks that caused yet again a renewed peak in monthly fatalities called for Israeli households to alter their perceived (subjective) level of personal risk and the possible effect on their future income. Note that it is also likely that terror attacks carried out by Jews against civilian and military Palestinians targets have caused Israeli households update their beliefs, as Israeli households expected revenge. However, appropriate data on such terror attacks was not available.

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[18] These are the most recently revised data series. National accounts data for 1980-1995 were compiled according to SNA68, whereas those for 1995 and onwards are based on SNA93.

Table 1: Total Number of Violent Attacks, Dead, and Wounded by Target of Attack 1980-2002

Category		Total			Israel			WBGS		
Major	Minor	Attacks	Dead	Wounded	Attacks	Dead	Wounded	Attacks	Dead	Wounded
1. **Armed Forces**	Military Personnel	219	286	505	40	47	110	179	239	395
	Police Personnel/Facility	12	8	10	4	3	4	8	5	6
2. **Government**	Government Personnel	2	2	1	1	1	-	1	1	1
3. **Civilians in General**	Civilians*	230	264	616	100	107	431	130	157	185
4. **Public Transportation**	Bus	44	246	1,018	28	212	934	16	34	84
	Bus stop	19	61	535	19	61	535	--	--	--
	Train Station	1	3	90	1	3	90	--	--	--
5. **Commercial Zone**	Shopping Center	14	53	986	13	52	986	1	1	-
	Marketplace	11	29	423	9	28	423	2	1	-
6. **Entertainment**	Entertainment Facility	5	42	236	5	42	236	--	--	--
	Restaurant	10	49	379	8	48	375	2	1	4
	Hotel	2	29	160	1	29	150	1	-	10
	Beach	3	1	-	3	1	-	--	--	--
7. **Vehicle**	Vehicle	69	83	97	8	11	48	61	72	49
	Cargo Transport	7	6	-	2	-	-	5	6	-
8. **Industrial Zone**	Industrial Zone	5	1	13	2	-	9	3	1	4
9. **Tourist**	Tourist	1	-	11	1	-	11	--	--	--
10. **Others**	Place of Worship	1	11	50	1	11	50	--	--	--
	Pupil	1	2	4	1	2	4	--	--	--
	School	1	-	8	1	-	8	--	--	--
	University	1	9	86	1	9	86	--	--	--
	Borders	6	23	42	5	20	42	1	3	-
Grand Total		664	1,208	5,270	254	687	4,532	410	521	738

* Including infiltrations, shootings, stabbings, and other events in different places.

Source: The database was compiled by the author from various sources listed in section 3.1.

Most of the attacks in Israel took place in Jerusalem and in a 10-mile-wide coastal strip that stretches along a 100 miles length from Haifa in the north to Ashdod in the south. During the period under investigation, over 80% of the Israelis, and only Israelis, lived or worked daily in this area. As Figure 2 shows,[19] terror attacks were primarily on civilian targets (accounting for more than 90% of the death toll).[20] The most lethal attacks targeted public transportation, commercial zones and entertainment facilities. Therefore, attacks inside Israel had direct implications on most Israeli households' personal security and future income.[21] Attacks against military or production targets in Israel numbered only a few (accounting for about 7% of dead).

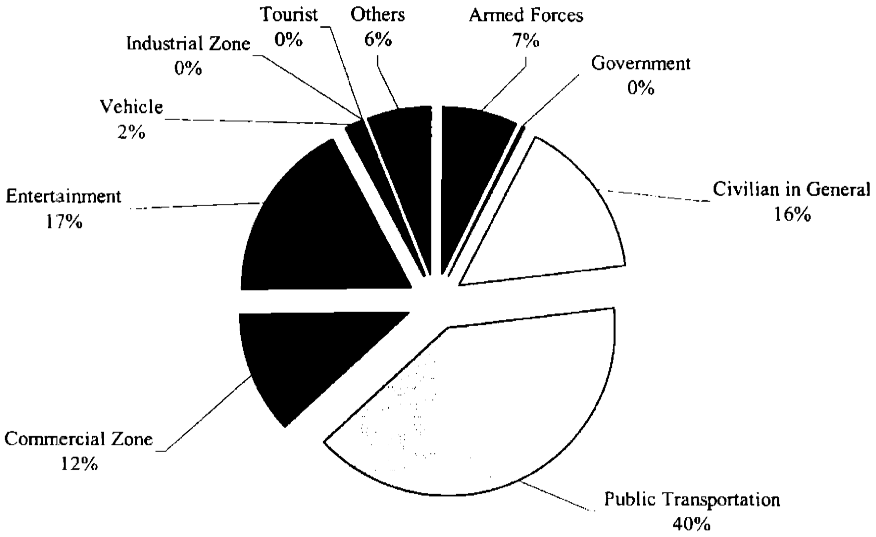

Figure 2: Total terror fatalities in Israel by target of attack, 1980-2002

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[19] The distribution of wounded provides basically a similar pattern.

[20] Different methods were used, of which shooting and grenades, suicide-carried explosives, and knife attacks were the most common as well the most deadly, accounting for more than 80% of all categories: number of attacks, number of dead, and number of wounded.

[21] One might argue that fluctuations in uncertainty fell more heavily on some households, while others (e.g., those who could afford to work in the countryside and use their private cars) were able to avoid it. If such were the case, then the impact of fluctuations in aggregate uncertainty on consumption was aggravated (see Blanchard and Mankiw, 1988).

Conversely, less than 3% of Israelis lived in settlements in the areas of the WBGS. A total population of approximately 200,000 lived in 150 Israeli settlements spread among nearly 700 Palestinian towns with a total population of just over three million Palestinians, as of March 2000.[22, 23] Both Israeli civilian and Israeli military targets were attacked in WBGS and equally suffered in the death toll (see Figure 3).

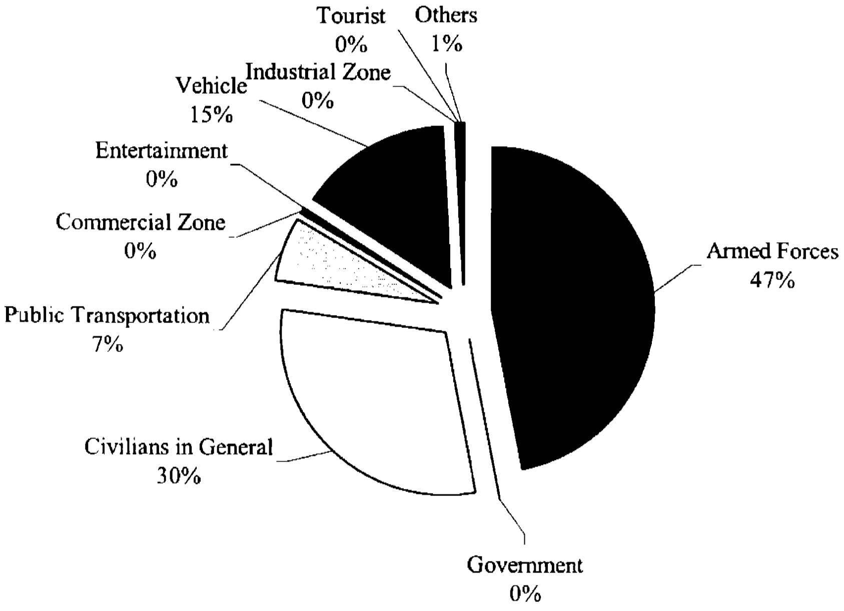

Figure 3: Total Terror Fatalities in WBGS by Target of Attack, 1980-2002

Notably, only the demand side (for consumption of durables, nondurables, and services) of the Israeli economy was targeted; industry, tourists, politicians, diplomats, and international interests

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[22] Settling WBGS by Israeli settlers started in 1967 and has never stopped. In addition to the above official settlements, there have always been tens of unofficial outposts and settlements, each with only a few settlers.

[23] These figures are based on data obtained from the Israeli Central Bureau of Statistics, the Palestinian Bureau of Statistics, and Peace Now Settlement Watch research.

were rarely targeted. The terror attacks therefore had a direct and first-order effect on the demand side of the economy. Furthermore, due to the above demographic and attack patterns, it is expected that Israeli household consumption demand would respond more substantially to news about fatal attacks in Israel than in WBGS; the former meant a more imminent increase in perceived personal risk and possible retaliation.

Israeli households and investors perceived terror to be temporary and they expected the Israeli government to constantly improve its effectiveness in counter-terror. Figure 1 supports these perceived beliefs, as it shows that there were periods when terror escalated, and, later on, were followed by de-escalation and calm periods.

Theory relating uncertainty to consumption, in conjunction with the economic consequences of the terror inflicted on the Israeli population, underlies the following four null hypotheses:

1. Null hypothesis 1: An increase in the number of terror fatalities has a negative impact on total consumption.
2. Null hypothesis 2: Durables are more affected by terror fatalities than nondurables.
3. Null hypothesis 3: Terror fatalities in Israel have a higher impact than those in WBGS.
4. Null hypothesis 4: A temporary increase in terror fatalities causes a bust-boom cycle in consumption, whereas a permanent increase in terror fatalities causes a one-time drop in consumption.

The next section will examine these null hypotheses. Before proceeding to the empirical results, two further points are worth noting. First, since the economy of the perpetrators of the terror had only marginal feedback on the Israeli economy, temporal variations in consumption of Israelis can be expected to have had little effect on the intensity of fatal attacks; therefore, endogeneity is unlikely to be a serious issue on a quarterly basis.

Second, the significant increase in terror fatalities that began in the year 2000 coincided with the end of the Hi-Tech boom. The peak in terror fatalities and the end of the Hi-Tech boom are independent exogenous influences to the economy that are known not to be uncorrelated. Furthermore, while the end of Hi-Tech occurred through a short period that lasted one or two quarters without further fluctuations, terror fatalities after the year 2000 continued to fluctuate for several quarters. Therefore, as noted by Eckstein and Tsiddon (2004), controlling for fluctuations in the Hi-Tech may reduce the negative values of the coefficients of terror fatalities and the marginal significance levels but would not reverse the basic results.[24]

3.3. The effect of terror on consumption

This section analyzes the impact of terror attacks on household consumption at the aggregate and disaggregated levels (see Figure 1). We first test and reject the null hypothesis that terror fatalities series is non-stationary. To test the above four hypotheses, we use quarterly data to regress the logarithmic difference of the real consumption (growth rate) on the logarithmic change in terror fatalities.[25] To avoid a logarithm of zero in some quarters, 1 is added to the terror fatalities series. To test for possible lag responses due to durability or market imperfection, we include contemporaneous and lagged regressors (until a lag coefficient is insignificant at 5% level). Possible auto-regression and moving average (ARMA) components are examined according to Box-Jenkins (1976) methodology with 5% significance level. We find in all cases that MA(1) or/and MA(2) should be included. The Newey-West (1987) estimator, which is robust for the remaining heteroskedasticity and autocorrelation of any type, is used and reported. Quarterly dummies are also included to de-seasonalize the data.

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[24] Eckstein and Tsiddon (2004) control for fluctuation in the Hi-Tech sector by including the first differences in logs of the NASDAQ index in real dollar terms for the period 1980:1 to 2003:3 as an additional exogenous variable in the VAR system with two and one lags.

[25] Another possibility is to regress the level of consumption on a time trend and the level of fatalities. On a macroeconomic theoretical basis, this should not change the results, and indeed it does not provide qualitatively different results. However, on an econometric theoretical basis, using the first differences as the dependent variables in a time series regression is preferable, because of their better treatment of serial correlation.

Null hypothesis 1: An increase in terror fatalities has a negative impact on total consumption

We test the first null hypothesis that terror fatalities have a negative impact on total consumption, with possible lagged response, by estimating the following regression:[26]

$$(1) \quad dLn(C_t) = \alpha + \sum_{i=0}^{4} \beta_i * dLn(TF_{t-i}) + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k}$$

where C_t is total consumption of Israelis, TF_t is the total number of Israeli terror fatalities in both Israel and WBGS, Qd_t is a quarterly dummy variable that equals 1 for quarter d and zero otherwise, and δ_k is the moving average component. The results of this regression are shown in column 1 in Table 2; for brevity, I only report the estimates of the constant and the coefficients of dLn(TF_{t-i})'s. The contemporaneous through the last significant lagged estimated coefficients of dLn(TF_{t-i}) are negative. Also, the one-year elasticity (i.e. long-term effect), which is the sum of all five coefficients of dLn(TF_{t-i}), is negative and significant. Therefore, we reject the null that fatal terror has no negative impact on total consumption. This impact starts after a delay of one quarter and lasts for not longer than a year, as only the first through the third lagged estimated coefficients are significant.

The estimated one-year impact of terror on consumption is substantial, as the one-year elasticity is about –0.023% (see last row first, column in Table 2), which means that a permanent 100% increase in quarterly fatalities (for instance, from 10 to 20) will bring about 2.3% decrease in total private consumption in the year following the increase in fatalities. Applying this estimate to the actual fatality numbers during 2000/Q1-2002/Q4 of *El-Aqsa Intifada*, the Palestinian uprising against the Israeli occupation, reveals a total contribution of roughly –9% to consumption growth rate. This finding, along with the fact that the average annual consumption growth rate was about 4% per annum for the five years preceding the *El-Aqsa Intifada*, explains the slightly negative (less than -1%) consumption growth rate during 2000/Q1-2002/Q4 of *El-Aqsa Intifada*, for the first time in two decades.

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[26] Note that an error correction model is not applicable in this case, since consumption is integrated of order one, i.e., I(1), while terror fatalities is stationary, i.e., I(0).

Table 2: The impact of terror fatalities on total consumption

	dLn(*Total consumption*)$_t$			
	1980/Q1- 2002/Q4	1980/Q1-2002/Q4 GARCH(1.1)	1980/Q1-2002/Q4 with IV for GDP	1987/Q1-2002/Q4
	(1)	(2)	(3)	(4)
C	**-0.052** *	**-0.057** *	**-0.053** *	**-0.054** *
	(0.00)	(0.00)	(0.00)	(0.00)
dLn(TF_t)	**-0.001**	**-0.003**	**0.000**	**-0.002**
	(0.76)	(0.46)	(0.86)	(0.21)
dLn(TF_{t-1})	**-0.008** *	**-0.007** **	**-0.007** **	**-0.004** *
	(0.04)	(0.06)	(0.07)	(0.04)
dLn(TF_{t-2})	**-0.008** *	**-0.005** **	**-0.008** *	**-0.004** **
	(0.01)	(0.09)	(0.03)	(0.06)
dLn(TF_{t-3})	**-0.007** *	**-0.006** *	**-0.007** **	**-0.005** **
	(0.05)	(0.04)	(0.09)	(0.07)
dLn(TF_{t-4})	**0.001**	**-0.001**	**0.001**	**0.001**
	(0.78)	(0.66)	(0.72)	(0.78)
R^2	0.72	0.72	0.71	0.84
D-W	1.97	1.97	1.97	2.07
#Obs.	87	87	85	66
One-year Elasticity	**-0.023** *	**-0.022** *	**-0.021** *	**-0.015** *
	(0.01)	**(0.03)**	**(0.05)**	**(0.05)**

Notes: One-year elasticity is the sum of all five coefficients of dLn(TF_{t-1}). Newey-West HAC Standard Errors & Covariance are used (lag truncation=3). *P*-values are in parenthesis; * and ** indicate variables significant at 5% and 10% significance level, respectively. To avoid a logarithm of zero in some quarters, 1 is added to *TF* series. All data employed were chained at 2000 prices from original series. Quarterly dummy de-seasonalizing variables and MA components are included but not reported for brevity. The instrumental variables in column (3) include 2-4 lagged GDP growth rates and 2-4 lagged consumption growth rates.

Different tests are used to examine the robustness of the above findings. The tests include GARCH specification and controlling for business cycles (by including contemporaneous

GDP).[27] The results of these tests, presented in columns (2)-(3) of Table 2, show no notable changes in the substance or significance of the coefficients. Further, I test for a possible breakpoint after the *first Intifada* (uprising) in 1987. T test is performed in two ways. First, regression (1) is run above for the period 1987-2002 rather than 1980-2002, yielding no notable changes in the results (compare results in column (4) with column (1) of Table 2). Second, the following regression is run for the entire period 1980-2002:

$$\begin{aligned} dLn(C_t) = \alpha + \sum_{i=0}^{4} \beta 80_i * (1 - D87_t) * dLn(TF_{t-i}) + \sum_{i=0}^{4} \beta 87_i * D87_t * dLn(TF_{t-i}) \\ + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k} \end{aligned} \tag{2}$$

where $D87_t$ is a dummy variable equals zero for the period 1980-1986, and one for the period 1987-2002. Next, we test the null hypothesis: $H_0 : \sum_{i=0}^{4} \beta 80_i = \sum_{i=0}^{4} \beta 87_i$, which is rejected at 5% percent significance level.

Other tests were also performed but are not reported in Table 2. To test the null hypothesis that consumption growth is affected by the level of, rather than the change in, terror fatalities, the fifth lag of terror fatalities (TF_{t-5}) was added as an additional regressor. If the coefficient of this regressor is significant, we cannot reject the null hypothesis that consumption growth rate should be regressed on contemporaneous and one-through-five lagged *TF* in levels; that is, we have co-integration. The results show that this additional regressor is insignificant for all cases, rejecting the null hypothesis.[28]

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[27] When GDP was included, I used instrumental variables estimation. The instrumental variables include 2, 3, and 4 lagged GDP growth rates and 2, 3, and 4 lagged consumption growth rates. We also run the same regression by using consumption-GDP ratio as dependent variable and find no change in the results.

[28] We also directly examine the above null by regressing the consumption growth rate on the level of terror fatalities (contemporaneous and five lags), and find that all coefficients of terror fatalities are insignificant, which rejects the null that a temporary terror attacks impact the long-run (steady-state) consumption level or growth rate.

To test the importance of the frequency of attacks, we include the number of attacks in regression (1) above and find it insignificant at 10% level. We also include quadratic and cubic values of TF_{t-i}'s and find them insignificant at 10% level.

Null hypothesis 2: Durables are more affected by terror fatalities than nondurables.

Next, we proceed to test our second hypothesis: the higher the durability of consumption, the greater the impact of terror fatalities is. To do so, we regress the following three regressions:

$$(3)\quad dLn(\mathrm{Dur}_t) = \alpha + \sum_{i=0}^{4} \beta_i * dLn(TF_{t-i}) + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k}$$

$$(4)\quad dLn(Non\mathrm{Dur}_t) = \alpha + \sum_{i=0}^{4} \beta_i * dLn(TF_{t-i}) + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k}$$

$$(5)\quad dLn(Ser_t) = \alpha + \sum_{i=0}^{4} \beta_i * dLn(TF_{t-i}) + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k}$$

where the variables Dur_t, $NonDur_t$, and Ser_t are durables consumption, nondurables consumption, and services consumption, respectively, in the domestic market. The results of these three regressions are presented in Table 3.

The results of the regressions are consistent with the theory. First, they provide similar qualitative results as those of regression (1); namely, terror fatalities negatively affect consumption of durables, nondurables, and services. This is evident from the finding that the contemporaneous through the last significant coefficients of *TF* as well as the long-term effects, measured by the one-year elasticities, are negative in all three regressions. Second, durables consumption seems to adjust slower than nondurables consumption. This is because the coefficient of one, two, and three lagged terror fatalities are significant in the durables regression, whereas it is only the coefficient of one lagged terror fatalities that is significant in the nondurables regression.

Table 3: The impact of terror fatalities on durables, nondurables, and services, 1980/Q1-2002/Q4

	dLn(*Durables consumption*)$_t$	dLn(*Nondurables consumption*)$_t$	dLn(*Services consumption*)$_t$
	(1)	(2)	(3)
C	**0.026**	**-0.058** *	**-0.029** *
	(0.38)	(0.00)	(0.00)
dLn(TF_t)	**-0.003**	**-0.001**	**-0.002**
	(0.77)	(0.84)	(0.33)
dLn(TF_{t-1})	**-0.032** *	**-0.008** *	**-0.004** *
	(0.04)	(0.04)	(0.05)
dLn(TF_{t-2})	**-0.037** *	**-0.005**	**-0.003** **
	(0.00)	(0.11)	(0.09)
dLn(TF_{t-3})	**-0.020** **	**-0.005**	**-0.004** *
	(0.07)	(0.27)	(0.02)
dLn(TF_{t-4})	**0.012**	**0.003**	**0.000**
	(0.37)	(0.52)	(0.92)
R^2	0.23	0.68	0.83
D-W	1.96	1.94	1.87
#Obs.	87	87	87
One-year Elasticity	**-0.079** *	**-0.015** *	**-0.012** *
	(0.02)	(0.02)	(0.00)

See notes on Table 2.

Third, the strongest long-term impact, measured by the one-year elasticity, is on durables consumption (-0.08), followed by nondurables consumption (-0.015) and services consumption (-0.012). In other words, a 100% increase in the number of fatalities (e.g., from 10 to 20 fatalities per quarter) will cause a drop by 8%, 1.5%, and 1.2% in durables, nondurables, and services consumption over the proceeding year, respectively.

The robustness of the above findings is examined by testing the impact of fatal attacks on the decomposition of durables consumption: furniture, household equipment, and personal transportation equipment.[29] We find that the qualitative results remain unchanged; namely, there is a negative impact on all of these three sub-categories with at least one-quarter delay. An examination of the sub-categories of the nondurables series shows that the food, beverages, and tobacco category (the least durable sub-category of the nondurables consumption) is hardly affected.

Although services consumption was the least affected as predicted by the theory, it is still interesting to understand why it is at all affected and the channel through which it is affected. To examine this matter, we consider the composition of the data series of services and find that services fluctuate mainly due to fluctuations in dining and accommodations services, which is substantially affected by foreign demand in the domestic market. Fleischer and Buccola (2002) found that foreign demand for accommodation at Israeli hotels is sensitive to terrorist activity, whereas domestic demand is insensitive. They also report that domestic demand provides only little buffer for declines in foreign tourism. Therefore, we expect that the elasticity of foreign consumption in the domestic market with respect to terror fatalities is higher than the elasticity of dining and accommodations services with respect to terror. I estimate both elasticities and present the results in Table 4. As expected, both elasticities are significantly negative and the elasticity of foreign consumption is higher than the elasticity of dining and accommodations services. Therefore, it appears that the impact of terror fatalities on services is primarily through the effect on foreign demand in the domestic market.

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[29] For brevity, the results are not reported and can be obtained from the author.

Table 4: The impact of terror fatalities on consumption- sub-categories, 1980/Q1- 2002/Q4

	dLn(*Foreigners' consumption in the domestic market*)	dLn(*Dining and accommodation services*)	dLn(*Non-profit institutions serving household*)
	1980/Q1-2002/Q4	1995/Q1-2002/Q4	1980/Q1-2002/Q4
	(1)	(2)	(3)
C	**-0.103** *	**-0.037** *	**0.038** *
	(0.00)	(0.03)	(0.00)
dLn(TF_t)	**-0.041** *	**-0.030** *	**0.008**
	(0.02)	(0.02)	(0.22)
dLn(TF_{t-1})	**-0.049** *	**-0.006**	**0.000**
	(0.01)	(0.57)	(0.93)
dLn(TF_{t-2})	**-0.026** *	**-0.014**	**-0.001**
	(0.05)	(0.16)	(0.82)
dLn(TF_{t-3})	**0.003**	**-0.018** *	**0.003**
	(0.82)	(0.01)	(0.43)
dLn(TF_{t-4})		**-0.003**	**0.001**
		(0.71)	(0.85)
R^2	0.68	0.87	0.34
D-W	2.16	2.48	1.88
#Obs.	88	30	87
One-year Elasticity	**-0.113** *	**-0.071** *	**0.010** *
	(0.00)	(0.00)	(0.01)

See notes on Table 2.

Note also that, unlike Israelis who respond in delay, foreign consumers adjust immediately. The coefficient of the contemporaneous terror fatalities is significantly negative in both regressions for foreign consumption in the domestic market as well as for dining and accommodation services. Further, not only is the timing of response different, but also the magnitude of the response: foreigners' elasticity (-0.113) is about five times that of Israelis (-0.023). Therefore, we should expect an increase in terror activity to have an immediate and intense negative effect on the outputs of all industries for which there is foreign demand, such as airlines, hotel accommodation, and restaurants.[30]

Finally, we examine the impact of fatal attacks on the consumption of non-profit institutions serving households (the last sub-category of consumption; see chart 1 above). We find that it is the only category of consumption that is positively impacted by fatal terror attacks, as its one-year elasticity is 0.01 (see Table 4, column 3). The rationale behind this exceptional positive sign is clear: when people are hit by terror attacks, more non-profit institutions will come to their aid.

Null hypothesis 3: Terror fatalities in Israel impact consumption more than those in WBGS

As noted above, perceived personal insecurity and uncertainty about future income is more sensitive to fatal attacks in Israel than in WBGS. Therefore, we expect that the impact of fatal terror attacks in Israel to be stronger than the impact of fatal attacks in WBGS. We test this null hypothesis by running regressions (3)-(5) using two alternative series: Israeli fatalities in Israel and Israeli fatalities in WBGS, instead of the sum of both. Since the impact of attacks against civilian targets may differ from the impact of attacks against military targets, and because attacks in WBGS are more military-target intense than attacks in Israel, we compare the results of regressions that employ fatalities in attacks against civilian targets in WBGS as against those of regressions that employ fatalities in attacks against civilian targets in Israel. The estimates of these regressions are presented in Table 5 below.

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[30] These results are not confined to the case when terror attacks involve the use of airplanes, as such was the case on September 11.

Table 5: The effect of terror fatalities on consumption by place of attack, 1980/Q1- 2002/Q4

	Attacks against civilians in Israel			Attacks against civilians in WBGS		
	dLn(*Durables consumption*)$_t$ (1)	dLn(*Nondurables consumption*)$_t$ (2)	dLn(*Services consumption*)$_t$ (3)	dLn(*Durables consumption*)$_t$ (4)	dLn(*Nondurables consumption*)$_t$ (5)	dLn(*Services consumption*)$_t$ (6)
C	**0.027**	**-0.057** *	**-0.029** *	**0.010**	**-0.063** *	**-0.030** *
	(0.42)	(0.00)	(0.00)	(0.75)	(0.00)	(0.00)
dLn(*TF*$_t$)	**-0.013**	**0.000**	**-0.003** *	**0.013**	**0.004**	**-0.002**
	(0.18)	(0.94)	(0.04)	(0.50)	(0.62)	(0.50)
dLn(*TF*$_{t-1}$)	**-0.021** *	**-0.010** *	**-0.005** *	**-0.032** *	**-0.002**	**-0.002**
	(0.05)	(0.01)	(0.03)	(0.03)	(0.75)	(0.25)
dLn(*TF*$_{t-2}$)	**-0.021** *	**-0.004**	**-0.002**	**-0.056** *	**-0.013** *	**-0.005** *
	(0.01)	(0.13)	(0.14)	(0.00)	(0.04)	(0.03)
dLn(*TF*$_{t-3}$)	**-0.015**	**-0.003**	**-0.003** *	**-0.009**	**-0.007**	**-0.003**
	(0.20)	(0.43)	(0.02)	(0.52)	(0.32)	(0.22)
dLn(*TF*$_{t-4}$)	**0.006**	**0.002**	**-0.001**	**0.023**	**0.005**	**0.000**
	(0.54)	(0.55)	(0.30)	(0.41)	(0.49)	(0.89)
R^2	0.17	0.68	0.82	0.21	0.6	0.82
D-W	1.93	1.90	1.85	1.98	1.94	1.88
#Obs.	87	87	87	87	87	87
One-year Elasticity	**-0.065** *	**-0.015** *	**-0.013** *	**-0.060** *	**-0.013** *	**-0.013** *
	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.00)

See notes on Table 2.

As expected, all significant coefficients and all one-year elasticities of consumption with respect to fatalities are negative, whether using fatalities in terror attacks against civilian targets in Israel or in WBGS, and whether the dependent variable is durables, nondurables, or services consumption. The one-year elasticities of durables and nondurables with respect to fatalities in attacks against civilian targets in Israel (-0.065 and -0.015) are not significantly greater than their counterpart elasticities with respect to fatalities in attacks against civilian targets in WBGS (-0.060 and –0.013). However, the one-year elasticity of services is equal for both fatalities in Israel and WBGS. A possible explanation for this equality is that foreigners, by whom this elasticity is most determined, do not distinguish between terror attacks that take place in Israel from those that take place in WBGS.

Null hypothesis 4: A temporary increase in terror fatalities causes a bust-boom cycle in durables consumption, whereas a permanent increase causes a one time drop

We simulate an impact on durables consumption in two scenarios: a temporary increase and a permanent increase in terror fatalities. We use the coefficients on Table 3 to report the results of these simulations in Figure 4 and Figure 5. The construction of our regression produces results that are consistent with the theory. The graphs of durables consumption show a bust-boom cycle and a one-time drop after, respectively, a temporary and a permanent increase in terror fatalities.

Finally, it is possible that there is asymmetry in the response of durables consumption to changes in terror fatalities: the impact of an increase is different than the impact of a decrease in terror fatalities. In Appendix A, we run simulations that allow asymmetry. These simulations reveal a similar pattern for the durables consumption as in Figure 4 and Figure 5, with two exceptions. First, the bust-boom cycle is stronger. Second, a temporary increase in terror causes durables consumption to retreat to a lower than the initial level, which suggests a stronger income shock effect (see Figure 6).

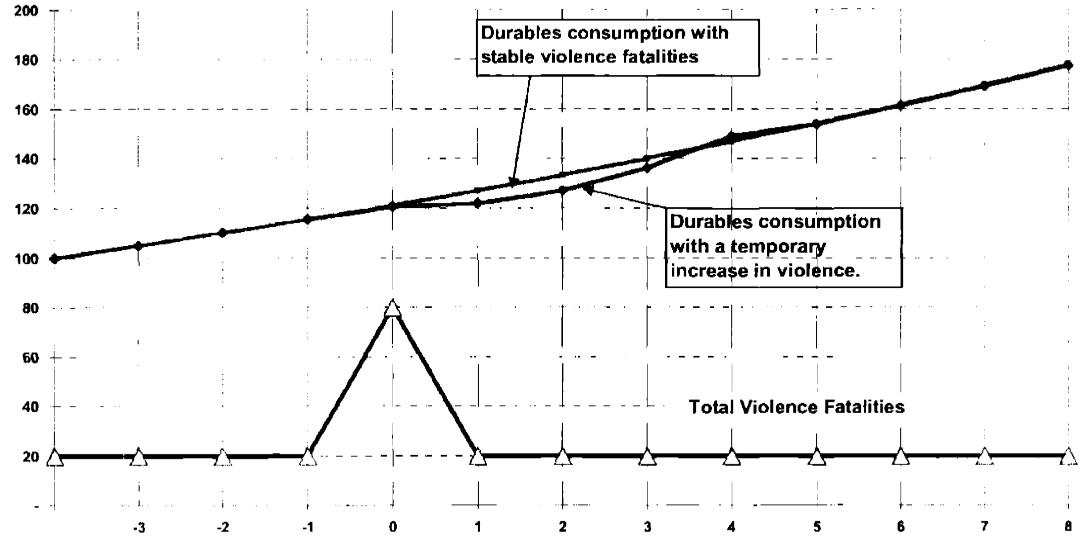

Figure 4: Simulation of a temporary increase in terror fatalities

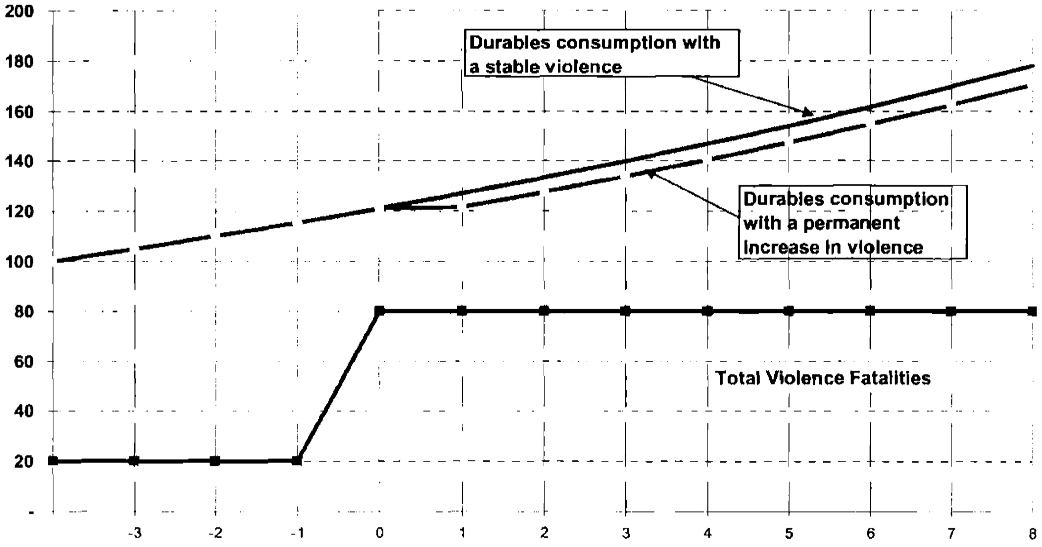

Figure 5: Simulation of a permanent increase in terror fatalities

3.4. The effect of terror on investment

This section examines how terror affects investment. To this end, the following regression is estimated:

$$(6)\quad dLn(FC_t) = \alpha + \sum_{i=0}^{4} \beta_i * dLn(TF_{t-i}) + \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k}$$

where FC_t is fixed capital and the regressors are the same as in regression (1). Also included are dummy variables to control for the 1991 Gulf war in regression (6). The results are presented in Table 6.

All the coefficients of terror fatalities as well as the one-year elasticity in regression (6) are negative. Lags one through three and the one-year elasticity are significant. This suggests that, as predicted by the irreversible investment theory discussed above, terror fatalities negatively impact irreversible investment through the effect on uncertainty. The one-year elasticity of fixed capital investment with respect to terror fatalities (-0.055) is higher than the one-year elasticity of total consumption with respect to terror fatalities (-0.023), which means that firms' decision on investment is more sensitive than households' decision on consumption.

Table 6: The impact of terror fatalities on investment, 1980/Q1-2002/Q4

	dLn(*Fixed capital*) 1980-2002	
	(1)	
C	**0.055**	*
	(0.00)	
dLn(TF_t)	**-0.005**	
	(0.45)	
dLn(TF_{t-1})	**-0.019**	*
	(0.01)	
dLn(TF_{t-2})	**-0.012**	*
	(0.03)	
dLn(TF_{t-3})	**-0.014**	*
	(0.00)	
dLn(TF_{t-4})	**-0.005**	
	(0.46)	
R^2	0.60	
D-W	2.06	
#Obs.	87	
One-year Elasticity	**-0.055**	
	(0.00)	

See notes on Table 2.

4. Summary and conclusions

This paper has demonstrated quite robust evidence for the effect of terror on consumption and investment. Uncertainty due to terror significantly affects consumption and investment in a manner consistent with the theory that relates uncertainty and consumption. A temporary increase in terror fatalities, while having no impact on the long-run level and growth rate of consumption, causes a bust-boom cycle in consumption in the short-run. Durables consumption is most affected, followed by nondurables. Services are affected primarily through the tourism channel, which affects dining and accommodation services. The responses of all series lag one quarter (except dining and accommodation services) and last no longer than three quarters. A permanent increase in terror fatalities causes a one-time drop in consumption.

Fixed capital also exhibits a bust-boom cycle after a temporary increase in terror fatalities and a one-time drop after a permanent increase in terror fatalities. However, the change in inventories is positively impacted by terror fatalities, which is an indication that the demand side reacts faster than the supply side.

Appendix A: Asymmetric response by consumption

In section 3 above, we assumed that the impact of an increase in terror fatalities has the same magnitude as the impact of a decrease in terror fatalities on durables consumption. Put another way, we assumed equal coefficients for both- positive and negative values of *TF*. We would like to test this hypothesis, while restricting ourselves to the case of durables. To this end, we run the following regression:

$$\begin{aligned} dLn(\mathrm{Dur}_t) &= \alpha + \sum_{i=0}^{3} \beta_i^P * P_{t-i} * dLn(TF_{t-i}) + \sum_{i=0}^{4} \beta_i^N * (1 - P_{t-i}) dLn(TF_{t-i}) \\ &+ \sum_{d=1}^{3} \chi_d Qd_t + \varepsilon_t + \sum_{k=1}^{K} \delta_k \varepsilon_{t-k} \end{aligned} \tag{7}$$

where P_{t-i} is one if there is an increase in terror fatalities at time t-i, and zero if there is a decrease in terror fatalities at time t-i. All other variables are defined as before. The estimates of this regression are presented in Table 7.[1]

As it is clear from Table 7, will still maintain the basic results that there is a negative correlation between terror fatalities and durables consumption and that the impact starts after one quarter and lasts no more than four quarters. Further, the elasticities of both- an increase and a decrease in terror fatalities- are significantly negative. Furthermore, we reject the null that both elasticities are equal at 5% significance level.

Simulations of the results of the above unrestricted regression for the case of temporary increase in terror fatalities are presented in Figure 6. The response of durables is similar to that in Figure 5 with two exceptions. First, the bust-boom cycle is stronger. Second, a temporary increase in terror causes durables consumption to retrieve to a lower level than the initial, which suggests a stronger income shock effect.

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[31] The number of lagged dependent variables was chosen such that until a lag coefficient is insignificant at 5% level and according to Box-Jenkins (1976) methodology.

Table 7: The impact of terror fatalities on durables consumption, assuming a different impact for a decease and an increase in terror fatalities, 1987/Q1-2002/Q4

	dLn(*Durables Consumption*)		
	(1)		
C	**0.042**		
	0.23		
P_t *dLn(TF_t)	**0.001**		$\sum_{0}^{3}\beta_i^P = -\underset{0.01}{0.070}$
	0.92		
P_{t-1} *dLn(TF_{t-1})	**-0.043**	*	
	0.05		
P_{t-2} *dLn(TF_{t-2})	**0.001**		
	0.96		
P_{t-3} *dLn(TF_{t-3})	**-0.029**	*	
	0.04		
$(1-P_t)$*dLn(TF_t)	**-0.010**		$\sum_{0}^{4}\beta_i^N = -\underset{0.05}{0.038}$
	0.62		
$(1-P_{t-1})$*dLn(TF_{t-1})	**0.023**		
	0.30		
$(1-P_{t-2})$*dLn(TF_{t-2})	**-0.071**	*	
	0.00		
$(1-P_{t-3})$*dLn(TF_{t-3})	**-0.022**		
	0.29		
$(1-P_{t-4})$*dLn(TF_{t-4})	**0.042**	*	
	0.04		
R^2	0.47		
D-W	2.08		
#Obs.	64		

See notes on Table 2. P_{t-i} is one if there is an increase in terror fatalities at time t-i, and zero if there is a decrease in terror fatalities at time t-i.

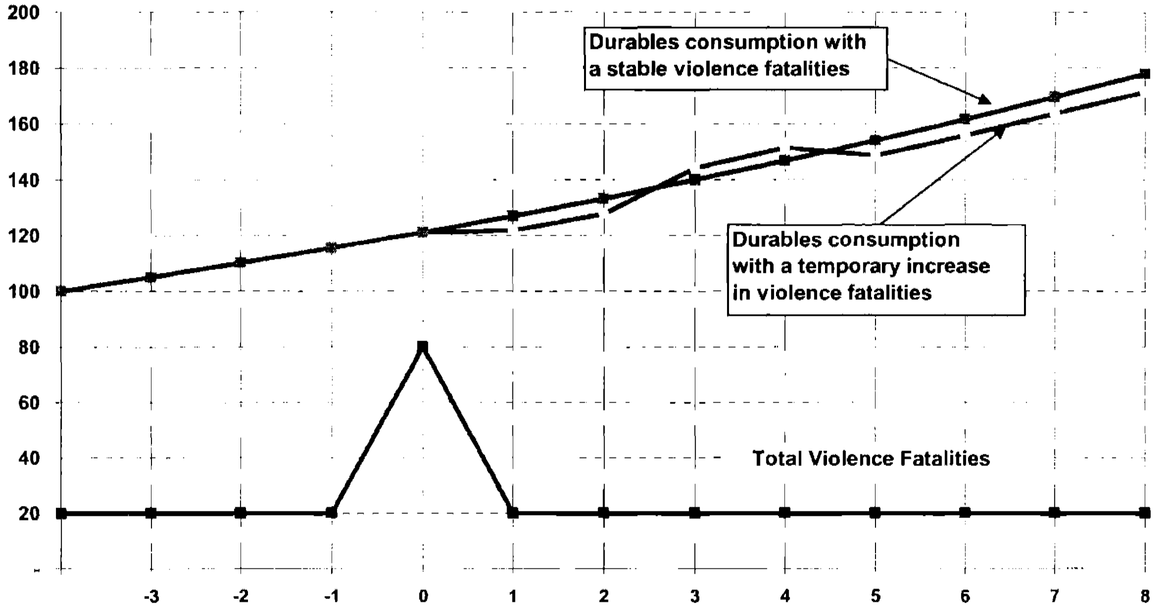

Figure 6: Simulation of temporary increase in terror fatalities with asymmetric response by durables consumption

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Chapter 2: Arbitrage Tests of Israel's Currency Options Markets

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Abstract

The aims of this study are threefold. First, we test the validity of the Black and Scholes (B-S) model as a naive option-pricing model for the case of an exchange-rate target zone. We find that although we cannot reject the weakly efficient market hypothesis (except for very-near-maturity deep-ITM options), we can reject the strongly efficient market and/or the B-S model validity hypotheses. The banking sector could have utilized arbitrage opportunities, notably for out-of-the-money, at-the-money, and far-from-maturity options, especially when employing inter-temporal weighted-average implied standard deviation.

Second, we estimate the liquidity premium for currency options by using a unique data set that allows us to comparing tradable and non-tradable options. The liquidity premium, though positive in average, is found to be negative for some options. This is an indication that there could have been arbitrage opportunities, especially for the banking sector. Third, we examine the null hypothesis that the Israeli currency options market is efficient, an issue that has not been investigated. Ex-post tests of arbitrage and dominance conditions do not permit rejection of the null hypothesis, except for very-near maturity, deep-in-the-money (ITM) options.

The paper enhances the literature by using a unique database from the Israeli currency options market, which includes currency options traded on the Tel Aviv Stock Exchange and (non-tradable) Bank of Israel currency options. In addition, this paper examines B-S when the exchange rate is confined to a target zone.

JEL classification: F31, G12, G13, G14.

Keywords: Currency Option, Black and Scholes, Liquidity Premium, Bank of Israel, Tel-Aviv Stock Exchange, Exchange Target Zone.

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1. Introduction*

This study focuses on the infant Israeli currency options market, which has not been thoroughly investigated so far. Currency options were first launched in Israel in 1987, by Bank HaMizrahi, but met with little success until 1989, when the Bank of Israel reversed its policy opposing such options. In November 1989 the Bank of Israel first launched currency options as part of its policy to encourage the financial and future instruments market. Currency options were first traded on the Tel Aviv Stock Exchange (TASE) in April 1994; over-the-counter (OTC) trading gained momentum only in the past year or so (1997–98).[1]

This study extends the existing literature on options in several ways. Most of the tests performed here have never before been applied to *currency* options. The study employs various methods to test the validity and efficiency of the currency options market and the validity of currency option models when the exchange rate is confined to a target zone.[2] To the best of my knowledge, this is the only study that compares OTC trading and the exchange in the same period for a single country. The study also evaluates the

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* This paper is part of my MA dissertation at the Hebrew University Business School and my Ph.D. thesis at MIT, Economics Department. I am indebted to Simon Benninga for his encouragement and valuable comments, Shlomo Yitzhaki for his patience, and late Maggie Eisenshtat for editorial help. I thank an anonymous referee for helpful suggestions and Hosny Zoaby and Erez Kapitolnik for excellent research assistance. I benefited from remarks by Gil Cohen, Ori Shaliv, Yona Shlomo, Tal Vardi, and participants of the econometric and financial lunches at MIT. I would also like to thank the Bank of Israel, the Tel Aviv Stock Exchange, and the Database at the Hebrew University for providing data and other helpful information. Finally, I thank Falk Institute for Economic Research in Israel for financial support. A previous version of this paper was published by Falk as Discussion Paper No. 99.01. All remaining errors are my own. A glossary of acronyms appears on p. 90.

[1] See Haj-Yehia (1997) for more on the development of the currency option market in Israel.

[2] See Haj-Yehia (1993) for an exposition and analysis of Israeli exchange rate regime.

liquidity premium by comparing the value of traded and non-traded — but otherwise identical — options.

The next section tests the null hypothesis that the Bank of Israel's currency options (BIO) and the Tel Aviv Stock Exchange options (TASEO) markets were efficient (a test of the "weak market-efficiency" hypothesis), by testing whether the options' prices were within the boundary arbitrage and dominance conditions. Section 3 examines the validity of the Black and Scholes model (Black and Scholes, 1973; henceforth referred to as B-S) as extended to the currency option. Given that the B-S model's assumptions are robust, this becomes a test of the "strong market-efficiency" hypothesis. Section 4 evaluates the liquidity premium, and surprisingly finds it to be negative for some options. The last section summarizes the main findings.

2. The Efficiency of the Currency Option Market

2.1. Introduction

Under the weakest rational assumption — that investors prefer having more to having less (non-satiated) — the absence of arbitrage opportunities constitutes a necessary, though not a sufficient condition for market efficiency. Since in a perfect market the existence of dominated securities would be equivalent to the existence of arbitrage opportunities,[3] I test the null hypothesis of efficient currency options markets by examining a set of boundary and dominance conditions for currency options prices (the weak market-efficiency hypothesis). These conditions, originally derived by Giddy (1983), Grabbe (1983), and Merton (1973), are summarized in Table 1; the results appear in Tables 2–4. However, some caveats before we proceed to the results are needed.

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[3] The existence of arbitrage opportunities means that with a zero investment position one can derive non-negative (not necessarily constant) returns in all future states of the world.

Table 1. Boundary Arbitrage Conditions for Rational Option Pricing [a]

a. Perfect markets		
a.1	The arbitrage upper boundary	$C(S,\tau,K) \le S$
a.2	The weak arbitrage lower boundary	$C(S,\tau,K) \ge Max(0,S-K)$
a.3	The strong arbitrage lower boundary	$C(S,\tau,K) \ge Max(0,Se^{-r^f\tau}-Ke^{-r^d\tau})$
b. Dominance conditions [b]		
b.1	Short \$ bond + long call $\succ$ Short NIS bond	$C \ge Max\left(0,(\frac{1}{T_s^{f^2} r_b^f}-T_a^c)S^r-\frac{K}{r_b^d}\right)$
b.2	Short NIS bond + short call $\succ$ Short \$ bond	$C \ge Max\left(0,(\frac{1}{T_s^{f^2} r_b^f}+T_a^c)S^r-\frac{K}{r_b^d}\right)$
b.3	Long NIS bond + long call $\succ$ Long \$ bond	$C \ge Max\left(0,(\frac{T_s^{f^2}}{r_l^f}-T_a^c)S^r-\frac{K}{r_l^d}\right)$
b.4	Long \$ bond + short call $\succ$ Long NIS bond	$C \ge Max\left(0,(\frac{T_s^{f^2}}{r_l^f}+T_a^c)S^r-\frac{K}{r_l^d}\right)$
c. The convexity condition		
c.1	The price convexity condition of K:	$\lambda C_1(S,\tau,K_1)+(1-\lambda)C_3(S,\tau,K_3) > C_2(S,\tau,K_2)$

a) See Appendix 1 for definitions of notations and parameters; NIS — New Israeli Shekel.

b) X $\succ$ Y means that X first-order-stochastically dominates Y. See derivation in Azolai and Landskroner (1993). Note that if one sector deviates from conditions (b.1) or (b.3), it would pay for that sector to sell more call options, whereas if another sector deviates from conditions (b.2) or (b.4), it would pay for that sector to buy more call options. Transactions occur when it pays one sector to sell call options at the same time it pays another sector to buy them.

2.2. *Caveats*

The first caveat has to do with the argument that testing for the existence of arbitrage opportunities requires an *ex-ante* test based on *transaction-by-transaction* data.[4] Such a test would be appropriate from the viewpoint of a trader who, upon observing a deviation, seeks to exploit it by submitting an order for the very next available

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[4] See, among others, Adams and Wyatt (1987), Bodurtha and Courtadon (1986), Galai (1977, 1978, and 1983), Shastri and Tandon (1985), Tucker (1985).

transaction, on the assumption that the price of the currency option will remain unchanged. However, the ex-ante nature of this strategy makes it risky, because prices (hence positive profits) are not known with certainty at the time the order is submitted. Moreover, a transaction-by-transaction database is not available for the TASEO, and only daily closing prices are available for BIO auctions. An ex-ante test based on the next day's closing prices is empirically questionable, and any reported deviations from such a test would be "mirages" rather than *bona fide* opportunities to make arbitrage profits. Therefore, we perform an *ex-post* test based on *daily* closing, high, and low prices to test for the existence of arbitrage opportunities.

The second caveat is that one should carefully consider whether a trader *can actually* exploit an observed arbitrage opportunity. The market might not be deep enough to allow him to pursue the chosen strategy, and it is therefore not clear whether or how many more contracts can be made if any above-normal profit opportunities present themselves. The depth of the market also influences transaction costs, so that small scale operations would find it difficult, if not impossible, to fill large orders (and even small orders take time to fill). This problem tends become more acute as time to maturity decreases and transaction costs increase. Furthermore, closing prices are sometimes artificially manipulated by market makers and therefore do not reflect the real prices that are sustained during the trading day or that regular investors can in fact realize for their transactions.

We examine the possibility of exploiting arbitrage opportunities by conducting an ex-post test that takes into account transaction costs and financial spreads.[5] This test is carried out for the two extreme cases of transaction costs — those borne by the banking sector and those borne by the household sector. The extreme case of high transaction costs is introduced to help explain observed deviations of actual closing prices from the

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[5] See Phillips and Smith (1980) for a classic exposition of the need for a careful analysis of market transaction costs and feasible trading rules in conducting option market efficiency tests. See also Black and Scholes (1973), Bodurtha and Courtadon (1986), Galai (1977), Gould and Galai (1974).

arbitrage boundary conditions. However, it should be remembered that market efficiency is determined with reference to the most efficient trader, that is, according to whether the banking sector can exploit deviations from the arbitrage boundary conditions in order to make above-normal riskless profits.

Finally, some of the variables are difficult to determine or observe. For example, the volatility of the exchange rate for the remaining lifetime of the option can only be *estimated* ex ante from historical data, with inconsistent results. In addition, an option should be matched with a daily foreign interest rate that has the same strike date, but the database of the London Inter-Bank Offered Rate (LIBOR) does not permit such matching; it provides monthly averages and is available only for specific discrete periods (one and seven days; one, three, six and 12 months' maturity). Here we use a proxy for the interest rates of different maturities — a linear interpolation based on two flanking rates.

2.3. The Results[6]

The results presented here in Tables 2–4 remain almost unchanged when using closing prices or high prices due to low volatility of prices during the day. The deviations decrease slightly as time passes and the market gains more experience, which suggests slow learning-by-doing.[7] Neither the BIO nor the TASEO have ever deviated from the arbitrage upper boundary condition. Deviations from the weak and strong arbitrage lower boundaries and the dominance conditions are detected only for the banking sector for very near (less than two weeks) maturities and for in-the-money (ITM) options (about 30% of the observations). Galai (1978) obtained similar findings for stock options in the Chicago Board Option Exchange (CBOE). Given that (a) interest rates, transaction costs (especially fixed costs), and the elasticities of the option price are all very sensitive to extremely short periods to maturity and to options being deep in the

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[6] A description of the data appears in Appendix 2.

[7] Not presented here for the sake of brevity.

money, and that (b) TASEO issues are not permitted during the option's last week, we conclude that the *weak market-efficiency* hypothesis cannot be rejected (see Tables 2–4).

The percentage of deviations by banks increases as the options draw closer to maturity or to deep ITM. The average deviation per contract remains almost unchanged (around NIS55 per contract for TASEO) in the case of closing prices, and decreases in the case of high prices. Comparing 2–4 months-to-maturity TASEO with 3-month BIO reveals that the average deviation and the percentage of deviations for BIO are much higher than those for TASEO. This does not mean that the Bank of Israel market is less efficient: TASEO and BIO have different depths, which affects the average and percentage of deviations.

The household sector did not deviate from the first two dominance conditions, but it did deviate considerably from the last two dominance conditions with regard to both the BIO and TASEO. Average deviations per contract and the percentage of deviations increase as TASEO tends towards deeper ITM or far-from-maturity, and decrease as BIO tends towards far-from-maturity. Differences in depth are the source of these differences in households' behavior as regards the BIO and the TASEO.

Note that the results of the convexity condition test applied by Haj-Yehia (1997) concur with the above results.

Table 2. Average and Number of Deviations from the Boundary Conditions:
BIO by Maturity [a]

	Time-to-maturity- τ (months)					
Boundary conditions	3-months Average	%	6-months Average	%	12-months Average	%
a.1	0	0	0	0	0	0
a.2	16	0	0	0	0	0
Closing price						
Banks						
a.3	84*	8	0	0	213	3
b.1	84*	8	0	0	213	3
b.2	84*	8	0	0	213	3
b.3	84*	8	0	0	238	3
b.4	84*	8	0	0	238	3
Households						
a.3	0	0	0	0	0	0
b.1	0	0	0	0	0	0
b.2	0	0	0	0	0	0
b.3	579*	99	400*	100	235*	53
b.4	712*	100	552*	100	298*	66
High price						
Banks						
a.3	118*	3	0	0	550	1
b.1	118*	3	0	0	550	1
b.2	118*	3	0	0	550	1
b.3	109*	3	0	0	560	1
b.4	109*	3	0	0	560	1
Households						
a.3	0	0	0	0	0	0
b.1	75	0	0	0	0	0
b.2	195	0	0	0	0	0
b.3	539*	99	382*	100	199*	45
b.4	671*	100	534*	100	285*	54
Sample size	944		84		138	

[a] Boundary condition labels follow Table 1. The sample includes the results of all the auctions held by the Bank of Israel between January 2, 1991, and May 2, 1996. The results are reported per 10 contracts, i.e., per $10,000 underlying asset (which makes them comparable with the results for TASEO). *t*-statistics are in parentheses. Asterisks indicate significant average deviations at the 5 percent level. The categorization of the BIO follows the Bank of Israel's categorization.

Table 3. Average and Number of Deviations from the Boundary Conditions:
TASEO by Maturity [a]

	Time-to-maturity- τ (months)									
Boundary conditions	0-.25		0.25-2		2-4		4-6		Total	
	Avg..	%	Avg..	%	Avg..	%	Avg..	%	Avg..	%
a.1	0	0	0	0	0	0	0	0	0	0
a.2	55*	29	62*	1	0	0	0	0	56*	1
Closing price										
Banks										
a.3	59*	27	62*	7	54*	3	59	1	59*	4
b.1	59*	27	62*	7	54*	3	59	1	59*	4
b.2	59*	27	62*	7	54*	3	59	1	59*	4
b.3	59*	27	63*	7	54*	3	61	1	60*	5
b.4	59*	27	63*	7	54*	3	61	1	60*	5
Households										
a.3	0	0	0	0	0	0	0	0	0	0
b.1	0	0	0	0	0	0	0	0	0	0
b.2	0	0	0	0	0	0	0	0	0	0
b.3	716*	71	562*	74	543*	74	567*	80	561*	75
b.4	859*	71	694*	76	677*	76	714*	80	697*	77
High price										
Banks										
a.3	51*	26	56*	6	61*	2	79	1	57*	4
b.1	51*	26	56*	6	61*	2	79	1	57*	4
b.2	51*	26	56*	6	61*	2	79	1	57*	4
b.3	42*	17	67*	3	48*	2	210	0	59*	2
b.4	42*	17	67*	3	48*	2	210	0	59*	2
Households										
a.3	0	0	0	0	0	0	0	0	0	0
b.1	0	0	0	0	0	0	0	0	0	0
b.2	0	0	0	0	0	0	0	0	0	0
b.3	707*	71	554*	74	533*	74	559*	80	552*	75
b.4	850*	71	686*	76	666*	76	706*	80	688*	77
Sample size	119		1,178		1,426		770		3,493	

[a] Boundary condition labels follow Table 1. The sample includes the results of all the auctions held by the Bank of Israel between January 2, 1991, and May 2, 1996. The results are reported per 10 contracts, i.e., per \$10,000 underlying asset (which makes them comparable with the results for TASEO). Standard deviations are in parentheses. Asterisks indicate significant average deviations at the 5 percent level. A TASEO is an ATM option if |S–K| < 0.05 NIS (New Israeli Shekels), an option is deep-ITM if K/S < 0.9. This choice is made in according to the way the options are issued on the issuing day.

Table 4. Average and Number of Deviations from the Boundary Conditions:
TASEO by Depth [a]

Boundary conditions	Depth (S/K)									
	OTM		ATM		ITM		Deep-ITM		Total	
	Avg..	%	Avg..	%	Avg..	%	Avg..	%	Avg..	%
a.1	0	0	0	0	0	0	0	0	0	0
a.2	0	0	36*	1	63*	12	0	0	56*	1
Closing price										
Banks										
a.3	0	0	36*	3	72*	33	113	19	59*	4
b.1	0	0	36*	3	72*	33	113	19	59*	4
b.2	0	0	36*	3	72*	33	113	19	59*	4
b.3	0	0	38*	3	72*	35	120	19	60*	5
b.4	0	0	38*	3	72*	35	120	19	60*	5
Households										
a.3	0	0	0	0	0	0	0	0	0	0
b.1	0	0	0	0	0	0	0	0	0	0
b.2	0	0	0	0	0	0	0	0	0	0
b.3	329*	69	651*	80	824*	75	831*	63	561*	75
b.4	452*	74	803*	80	978*	75	981*	63	697*	77
High price										
Banks										
a.3	0	0	36*	2	66*	31	120	19	57*	4
b.1	0	0	36*	2	66*	31	120	19	57*	4
b.2	0	0	36*	2	66*	31	120	19	57*	4
b.3	0	0	42*	1	62*	22	103	19	59*	2
b.4	0	0	42*	1	62*	22	103	19	59*	2
Households										
a.3	0	0	0	0	0	0	0	0	0	0
b.1	0	0	0	0	0	0	0	0	0	0
b.2	0	0	0	0	0	0	0	0	0	0
b.3	323*	69	640*	80	815*	75	834*	63	552*	75
b.4	446*	74	792*	80	969*	75	834*	63	688*	77
Sample size	1,246		1,955		276		16		3,493	

[a] See note for Table 3.

3. Validating the Options Pricing Model

3.1. The Model

In order to uniquely determine options pricing, the rational options pricing model (OPM) adds structural assumptions to the aforementioned set of boundary conditions. These assumptions relate to the evolution of the underlying asset's price, the domestic interest rate, and the foreign interest rate. Therefore, if deviations of the boundary conditions are found for a given data set, they will reappear when any specific OPM is tested using the same data. The market cannot be shown to be both inefficient under weak conditions and at the same time efficient for compatible but stronger assumptions.

Based on the Black and Scholes (1973) and Merton (1973) valuation formula, Biger and Hull (1983) and Garman and Kohlhagen (1983) were the first to derive a currency options pricing model.[8] Their formula (henceforth: the B-S formula or model) is:[9]

$$C = e^{-r^f \tau} SN\left(\frac{\ln(S/K) + (r^d - r^f + \sigma^2/2)\tau}{\sigma\sqrt{\tau}} \right) - e^{-r^d \tau} KN\left(\frac{\ln(S/K) + (r^d - r^f - \sigma^2/2)\tau}{\sigma\sqrt{\tau}} \right)$$

Following Haj-Yehia (1997), this formula is used here as a naive formula for the Israeli market.

The empirical tests of B-S model validity are tests of joint null hypotheses: (i) the robustness of the assumptions; and (ii) market efficiency. In Haj-Yehia (1997) I tested the first null hypothesis and found that the B-S assumptions are fairly robust in the Israeli currency option market, except for very-near-maturity and deep ITM currency options. Therefore, assuming that these findings are correct, the tests of the B-S model's validity are tests of the null hypothesis that the market is efficient under the assumptions of that model (the "strong market-efficiency" hypothesis).

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[8] Stulz (1982) also developed a series of analytical formulas for European call and put options on the minimum or maximum of two risky assets; these, too, can be applied to value options.

[9] See Appendix 1 for definitions of notations and parameters.

3.2. The validity of the B-S model

If the B-S model is valid, i.e., if the B-S model's assumptions are robust and the currency option market is efficient, then:

(a) Implied standard deviation (ISD) should equal actual standard deviation (ASD).[10]

(b) Model price should equal actual price.[11]

(c) Dynamic neutral hedging should yield a risk-free rate of return.[12]

This subsection tests the validity of the B-S model, by examining whether the above predictions hold.[13]

(a) Comparing ISD with ASD

ISD is often said to be forward-looking and therefore a better estimator of volatility than one based on historical data. If the B-S formula is indeed correct and the market is efficient, then ISD should recover future volatility of the exchange rate without error, it should be numerically identical to the future volatility regardless of the set of options used (that have the same expiration date), and it should be stable in time (since the B-S model assumes constant volatility). In fact, however, this never occurs, a finding which is examined and found in this section as well. Therefore, we proceed by comparing the ISD with other standard deviation estimators in search for the estimator that is most likely to be substituted in the B-S formula by market participants when pricing currency options in Israel.

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[10] See Blomeyer and Klemkosky (1982), MacBeth and Merville (1979, 1980), and Thorp and Gelbaum (1980).

[11] See Goodman, Ross and Scmidth (1985), Tucker (1985), and Shastri and Tandon (1986).

[12] See Bodurtha and Courtadon (1986), Goodman (1985), Ng (1989), Shastri and Tandon (1986), Tucker (1985), and Tucker and Peterson (1988).

[13] Note that although the first and second tests presented in sub-sections (i) and (ii) can be performed for both options markets (Bank of Israel and the TASE), their results cannot be categorized by maturity for BIO because BIO are not traded.

In this study, for each trading day (t) I employ closing prices to calculate the following estimators of the volatility:

(a) The non-adjusted ISD of each option separately.

(b) The lag-non-adjusted ISD of each option separately (LISD), the inter-temporal weighted average ISD from the previous 7 trading days of the same option (IWISD), the arithmetical weighted average of the ISD from the previous day for all options traded at time ¯t1 that have the same maturity date (AWISD).

(c) The ASD of each option separately.

(d) The SD for a window of 7 days around t (WSD7D).

(e) The SD for 7 days, 2 months, 4 months and 6 months, historical data (HSD7D, HSD2M, HSD4M and HSD6M, respectively).

Next I calculate the deviation of the ISD from the ASD, WSD7D, and HSD. The deviation of the ISD from the ASD tests the validity of the B-S model and its ability to predict the future volatility or the efficiency of the market in predicting the ASD. The deviation of the ISD from the WSD7D tests whether the ISD is related to anticipated events that increase short-run volatility. Deviations of the ISD from the HSD test whether traders actually make use of historical data when calculating the SD to be substituted in the B-S formula.

Since only closing prices are used, one might argue that the deviations might have been caused by non-synchronization, or that the markets are not in equilibrium. To address this problem, we also calculate the number of times the ASD deviates from the non-adjusted ISD derived from the highest and lowest prices of each option on each trading day. These ISDs should be boundaries for the ASD set by the B-S model.

This section also calculates the deviation of the ISD derived from the Bank of Israel's minimal premium that was allowed to be submitted in the BIO's auctions. From this figure we can learn precisely how the Bank of Israel predicte the ASD, or by how much the Bank of Israel was willing to subsidize its currency options. All the results pertaining to this section are given in Tables 5–7.[14]

Table 5 shows a marked feature of the options: the market exhibited non-rational behavior by systematically underestimating the ASD. In other words, the B-S model fails to predict the ASD; which means rejection of the currency option market strong efficiency and the B-S model validity hypotheses. The ISD of the 3-month and 6-month BIO was underestimated, but not significantly different from zero when using WSD7D or HSD7D, and the under-estimation increases the longer the time horizon used for calculating the HSD. This means that market participants were more influenced by recent news and developments than by the more remote history when they calculated the SD to be substituted in the B-S formula.

The Bank of Israel also underestimates the ASD. A reasonable explanation is its desire to market these options and attract more participants, though this is a less reasonable explanation in the case of the 12-month BIO. The ISD is significantly overestimated for the 12-month option and the deviations are always larger, in absolute terms, than for the 3- and 6-month options. Since the closing premium of the 12-month option was always equal to the minimal premium determined by the Bank of Israel, the over-determined premium must have been a primary reason for the failure of the Bank of Israel to market the 12-month option.

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[14] The drawback of this approach is the fundamental inconsistency in using the B-S model to obtain predictions of a presumably nonstationarity variance, or in using weighted averages to estimate the "appropriate" ISD.

Table 5. Average Deviation of ISD from ASD and HSD:
BIO by Maturity and Depth [a]

	3-months			6-months	12-months
	ITM	OTM	Total	OTM	OTM
$\sigma_I - \sigma_A$	-0.0074*	-0.0105*	-0.0095*	-0.0182*	0.0166*
	(-8.85)	(-5.94)	(-7.94)	(-8.27)	(5.7)
$\sigma_I - \sigma_{W7D}$	-0.0007	-0.0002	-0.0004	-0.0008	0.0208*
	(-0.62)	(-0.13)	(-0.32)	(-0.33)	(5.46)
$\sigma_I - \sigma_{H7D}$	-0.0002	-0.0006	-0.0005	0.0016	0.0189*
	(-0.17)	(-0.3)	(-.335)	(0.65)	(4.88)
$\sigma_I - \sigma_{H2M}$	-0.0047*	-0.0052*	-0.0051*	-0.0022	0.0128*
	(-6.6)	(-3.18)	(-4.6)	(-1.73)	(4.25)
$\sigma_I - \sigma_{H4M}$	-0.0066*	-0.0062*	-0.0063*	-0.0029*	0.0110*
	(-9.41)	(-4.02)	(-6.14)	(-3.15)	(3.92)
$\sigma_I - \sigma_{H6M}$	-0.0074*	-0.0071*	-0.0072*	-0.0002*	0.0097*
	(-11.1)	(-4.65)	(-7.08)	(-3.24)	(3.37)
$\sigma_{MBI} - \sigma_A$	-0.0085*	-0.0139*	-0.0120*	-0.0227*	0.0170*
	(-7.16)	(-6.18)	(-6.71)	(-8.04)	(7.15)
Sample size	330	614	944	84	138

[a] See note to Table 2.

Table 6- Table 7 show that the B-S model's inability to predict the ASD does not depend on the depth or maturity of the option. However, the ISD is found to be higher than the HSD, and the gap widens as the options approach maturity. Our conclusion is that option traders use more information than merely historical data in determining the future volatilities that they substitute in the B-S formula.[15] The traders determine the ISD as a weighted average of the HSD and their expected ASD; less weight is assigned to the HSD as the option approaches maturity. The bottom parts of Tables 6 and 7 show that the inaccurate prediction of the ASD by the B-S model is not due to non synchronization or inequilibrium of the markets, regardless of the choice of ISD estimators. (For the sake of brevity, the tables contain only three estimators).

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[15] This is consistent with Chiras and Manaster (1978) and other studies cited there.

Table 6. Average Deviation of ISD from ASD and HSD:
TASEO by Depth [a]

	Depth (S/K)				
	OTM	ATM	ITM	Deep-ITM	Total
σ_I	0.0613	0.0530	0.0297	0.2427	0.05
$\sigma_I - \sigma_{LI}$	0.0011*	0.0001	-0.0024	0.0364	0.0004
	(4.9)	(0.06)	(-0.7)	(0.7)	(0.848)
$\sigma_{IWI} - \sigma_A$	-0.0090*	-0.0224*	-0.0869*	0.1862*	-0.0212*
	(-10.5)	(-19.5)	(-10.4)	(2.9)	(-19.71)
$\sigma_{AWI} - \sigma_A$	-0.0072*	-0.0212*	-0.0870*	0.1064	-0.0206*
	(-8.3)	(-17.2)	(-10.5)	(1.4)	(-18.55)
$\sigma_I - \sigma_A$	-0.0056*	-0.0218*	-0.0947*	0.1951*	-0.0206*
	(-7.3)	(-17.7)	(-10.9)	(2.3)	(-17.71)
$\sigma_I - \sigma_{W7D}$	0.0181*	0.0095*	-0.0211*	0.2059*	0.0111*
	(26.1)	(12.6)	(-5.8)	(2.4)	(15.06)
$\sigma_I - \sigma_{H7D}$	0.0190*	0.0102*	-0.0169*	0.2050*	0.0121*
	(27.0)	(14.3)	(-5.3)	(2.4)	(17.09)
$\sigma_I - \sigma_{H2M}$	0.0158*	0.0067*	-0.0219*	0.2018*	0.0086*
	(32.5)	(11.4)	(-7.4)	(2.4)	(13.4)
$\sigma_I - \sigma_{H4M}$	0.0140*	0.0064*	-0.0186*	0.2039*	0.0080*
	(31.6)	(11.4)	(-6.5)	(2.4)	(12.94)
$\sigma_I - \sigma_{H6M}$	0.0151*	0.0072*	-0.0174*	0.2060*	0.0090*
	(35.8)	(13.1)	(-6.1)	(2.5)	(14.63)
Percentage of cases in which ASD is within the ISDs implied by high and low prices					
$\sigma_I^l < \sigma_A < \sigma_I^h$	5.9%	7.5%	6.5%	12.5%	6.9%
$\sigma_{IWI}^l < \sigma_A < \sigma_{IWI}^h$	6.1%	8.4%	4.7%	12.5%	7.3%
Sample size	1,246	1,955	276	16	3,493

[a] See note for Table 3.

Table 7. Average Deviation of ISD from ASD and HSD:
TASEO by Maturity[a]

	Time-to-maturity - τ (months)				
	0-.25	0.25-2	2-4	4-6	Total
σ_I	0.0600	0.0585	0.0528	0.0528	0.0550
$\sigma_I - \sigma_{LI}$	0.0030* (4.92)	0.0009 (.06)	-0.0002 (-.74)	0.0003 (.69)	0.0004 (.848)
$\sigma_{IWI} - \sigma_A$	-0.1377* (-10.5)	-0.035* (-19.5)	-0.0104* (-10.38)	-0.0052* (2.87)	-0.0212* (-19.71)
$\sigma_{AWI} - \sigma_A$	-0.1339* (-8.27)	-0.0333* (-17.2)	-0.0111* (-10.5)	-0.0044* (1.36)	-0.0206* (-18.55)
$\sigma_I - \sigma_A$	-0.1339* (-7.28)	-0.0333* (-17.66)	-0.0111* (-10.93)	-0.0044* (2.27)	-0.0206* (-17.71)
$\sigma_I - \sigma_{W7D}$	0.0065* (26.06)	0.0146* (12.65)	0.0089* (-5.8)	0.0105* (2.42)	0.0111* (15.06)
$\sigma_I - \sigma_{H7D}$	0.0170* (27.01)	0.0146* (14.34)	0.0102* (-5.27)	0.0110* (2.41)	0.0121* (17.09)
$\sigma_I - \sigma_{H2M}$	0.0125* (32.54)	0.0109* (11.35)	0.0068* (-7.42)	0.0077* (2.42)	0.0086 (13.4)
$\sigma_I - \sigma_{H4M}$	0.0129* (31.63)	0.0111* (11.36)	0.0061* (-6.53)	0.0059* (2.44)	0.0080* (12.94)
$\sigma_I - \sigma_{H6M}$	0.0131* (35.79)	0.0120* (13.08)	0.0069* (-6.07)	0.0076* (2.47)	0.0090* (14.63)
Percentage of cases in which ASD is within the ISDs implied by high and low prices					
$\sigma_I^l < \sigma_A < \sigma_I^h$	3.4%	5.6%	8.8%	5.7%	6.9%
$\sigma_{IWI}^l < \sigma_A < \sigma_{IWI}^h$	0.0%	5.9%	9.5%	6.4%	7.3%
$\sigma_{AWI}^l < \sigma_A < \sigma_{AWI}^h$	5.9%	5.4%	8.6%	9.6%	7.6%
Sample size	119	1,178	1,426	770	3,493

[a] See note for Table 3.

The ISD decreases very slowly as the TASEO draws closer to maturity, but drops steeply as the TASEO goes further into deep-ITM. That is why the ISD of TASEO is higher than the HSD for ATM and ITM but lower than the HSD for ITM options.

We also learn from Tables 6–7 that the LISD is the best predictor of itself, except for very near-to-maturity options, and does not depend on depth.

(b) Model prices versus actual prices

If the model is valid and markets are efficient and synchronized (i.e., even if the ISD deviates from the other estimated SDs), the model's prices should not differ much from actual prices. For each trading date, t, we compute the model's price using the different estimators for volatility suggested in the previous section. We first test the significance of the average deviation between model and actual prices (see Tables 8–10).[16]

It has been argued (Galai, 1983) that comparing actual prices with model prices should yield the same results as a comparison between ISD and estimated SD. Tables 8–10 demonstrate that this is not necessarily so: overestimation of the ISD on *average* does not mean overestimation of the option price on *average*.

The signs and significance levels in Tables 8–10 are consistent with those in Tables 5–7 except for three cases: OTM 3-month BIO, very close to maturity TASEO, and LISD. This inconsistency can be explained precisely by the statement made in the previous paragraph, and becomes crucial in extreme cases (such as being very close to maturity), when a small deviation in SD leads to a large deviation in the option's price.

Tables 8–9 show that model prices are significantly different from actual prices, leading to rejection of the currency option market strong-efficiency and the B-S model validity

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[16] The basic results do not change by converting the absolute figures into relative deviations, provided neither the model price nor the actual price is very close to zero (which may bias the estimations). However, relative deviations become useless when the discussion turns to excess hedge returns, because the latter cannot be converted into percentages (the strategies are designed to be self-financed), and they are therefore not presented here.

hypotheses, which is consistent with the rejection in subsection (i). The negative average deviation of the actual option price from the model price based on the ASD is consistent with the underestimation of the ISD relative to the ASD. The underpricing of the options increases as the option draws closer to maturity and goes deeper ITM. The bottom parts of Tables 8–10 provide another facet of the systematic error in predicting the ASD, suggesting that this systematic error is not due to non-synchronization or to market inequilibrium, regardless of the estimation method of the ISD. This is inconsistent with the findings and conclusions of Black and Scholes (1972, p. 405), who state that (for stock options) "if the B-S model has an accurate estimation of the ASD, it works very well."

Tables 8–9 show that OTM options were overpriced (except when applying ASD), and ITM options were underpriced (deeply undepriced when applying ASD). These results are similar to those obtained by Blomeyer and Klemkosky (1982) for stock options, and to Black (1975) and Thorp and Gelbaum (1980), who maintain that their experience in trading CBOE options applying the B-S model is that this model tends to underprice OTM options while correctly pricing ITM and ATM options. Our results, however, contradict those obtained by MacBeth and Merville (1979), who observe that the deviations of the actual price from the B-S model price tend to be positive (negative) constant for deep-ITM (OTM).

As noted by Campbell, Lo and McKinlay (1997, ch. 9.3.3) the accuracy of the estimated option price depends on the strike price, the exchange rate, and the time to maturity. In some — though by no means all — cases, as time to maturity declines one may observe more biases of the model and actual prices. Other sources of bias are the presence of stochastic volatility or a misspecification of the bonds and exchange rate dynamics.

Again we observe that the Bank of Israel underestimated the 3-month and 6-month BIO, and overestimated the 12-month BIO. The underestimation of the 6-month BIO was greater than the 3-month BIO underestimate, possibly owing to the Bank's concern that marketing this new type of BIO series might fail.

Although the LISD is shown to be a good predictor of the ISD, the model price based on the LISD is underestimated, but not by much.

Table 8. Average Deviation of Model Price from Actual Price:
BIO by Maturity and Depth [a]

	3-months			6-months	12-months
	ITM	OTM	Total	OTM	OTM
$C^{a}-C^{m}_{A}$	-0.26*	-0.14*	-0.28*	-151*	192*
	(-14.09)	(-12.09)	(-17.65)	(-8.17)	(11.32)
$C^{a}-C^{m}_{W7D}$	-14	-43*	-24*	7	176*
	(-1.34)	(-8.67)	(-5.61)	(0.33)	(9.32)
$C^{a}-C^{m}_{H7D}$	16	-3	10*	13	182*
	(1.50)	(-0.46)	(1.95)	(0.62)	(7.2)
$C^{a}-C^{m}_{H2M}$	15*	-0.1	10*	-17*	166*
	(2.26)	(-0.02)	(1.98)	(-1.67)	(6.36)
$C^{a}-C^{m}_{H4M}$	7	-26*	-5	-23*	133*
	(1.22)	(-6.44)	(-1.25)	(-3.07)	(6.43)
$C^{a}-C^{m}_{H6M}$	4	-36*	-10*	-20*	122*
	(0.79)	(-9.38)	(-2.8)	(-3.54)	(6.14)
$C_{MBI}-C^{m}_{A}$	1	-41*	-13*	-187*	109*
	(0.27)	(-11.1)	(-3.83)	(-8.00)	(5.60)
$\lvert C^{a}-C^{m}_{A}\rvert$	16	8	13	16	20
$\lvert C^{a}-C^{m}_{W7D}\rvert$	11	7	9	14	25
$\lvert C^{a}-C^{m}_{H7D}\rvert$	12	9	11	14	28
$\lvert C^{a}-C^{m}_{H2M}\rvert$	13	9	11	8	28
$\lvert C^{a}-C^{m}_{H4M}\rvert$	10	6	9	6	22
$\lvert C^{a}-C^{m}_{H6M}\rvert$	9	6	8	4	21
Sample size	614	330	944	84	138

[a] See note to Table 2.

Table 9. The average deviation of model price from actual price:
TASEO by Depth[a]

	Depth (S/K)				
	OTM	ATM	ITM	Deep-ITM	Total
$C^a - C^m_{LI}$	3*	-3*	-52*	-32	-5*
	(3.2)	(-2.5)	(-10.6)	(-1.3)	(-5.2)
$C^a - C^m_{IWI}$	11*	6*	-48*	9	3*
	(9.5)	(3.9)	(-9.4)	(0.3)	(3.2)
$C^a - C^m_{AWI}$	3	-1	-42*	23	-3*
	(1.8)	(-0.5)	(-8.2)	(0.7)	(-1.6)
$C^a - C^m_{A}$	-31*	-61*	-127*	23	-55*
	(-8.2)	(-16.5)	(-11.2)	(0.7)	(-20.7)
$C^a - C^m_{W7D}$	65*	35*	-48*	23	39*
	(19.6)	(13.5)	(-6.7)	(0.7)	(19.4)
$C^a - C^m_{H7D}$	68*	36*	-43*	23	41*
	(20.8)	(15.2)	(-6.9)	(0.7)	(21.7)
$C^a - C^m_{H2M}$	59*	35*	-37*	23	38*
	(27.1)	(18.6)	(-6.1)	(0.7)	(25.9)
$C^a - C^m_{H4M}$	49*	35*	-34*	23	34*
	(28.6)	(20.0)	(-5.7)	(0.7)	(26.4)
$C^a - C^m_{H6M}$	55*	38*	-33*	23	39*
	(35.2)	(22.7)	(-5.6)	(0.7)	(30.5)
$\left\|C^a - C^m_{LI}\right\|$	25	43	70	80	39
$\left\|C^a - C^m_{IWI}\right\|$	29	45	71	102	42
$\left\|C^a - C^m_{AWI}\right\|$	44	50	70	109	50
$\left\|C^a - C^m_{A}\right\|$	83	116	158	108	108
$\left\|C^a - C^m_{W7D}\right\|$	105	90	88	109	95
$\left\|C^a - C^m_{H7D}\right\|$	106	85	79	109	92
$\left\|C^a - C^m_{H2M}\right\|$	74	66	76	109	70
$\left\|C^a - C^m_{H4M}\right\|$	58	62	75	109	62
$\left\|C^a - C^m_{H6M}\right\|$	61	63	75	109	63
Percentage of cases in which C^a does not deviate from boundaries					
$C^l_t < C_A < C^h_t$	15.2%	14.1%	2.9%	12.5%	13.6%
$C^l_{IWI} < C^a < C^h_{IWI}$	15.7%	10.0%	0.4%	0.0%	11.3%
Sample size	1,246	1,955	276	16	3,493

[a] See note for Table 3.

Table 10. The average deviation of model price from actual price:
TASEO by Maturity [a]

	Time-to-maturity - τ (months)				
	0-0.25	0.25-2	2-4	4-6	Total
$C^a - C_{LI}^m$	-30*	-4*	-5*	-3	-5*
	(-4.9)	(-3.0)	(-3.1)	(-1.2)	(-5.2)
$C^a - C_{IWI}^m$	-24*	3*	3*	8*	3*
	(-3.9)	(2.1)	(2.0)	(3.1)	(3.2)
$C^a - C_{AWI}^m$	-28*	-3	1	-5	-3
	(-3.9)	(-1.4)	(0.4)	(-1.4)	(-1.6)
$C^a - C_A^m$	-159*	-83*	-43*	-21*	-55*
	(-7.9)	(-17.4)	(-11.7)	(-3.7)	(-20.7)
$C^a - C_{W7D}^m$	-21*	30*	43*	54*	39*
	(-3.2)	(12.1)	(12.7)	(10.5)	(19.4)
$C^a - C_{H7D}^m$	-18*	30*	48*	55*	41*
	(-2.8)	(12.3)	(14.8)	(11.7)	(21.7)
$C^a - C_{H2M}^m$	-19*	29*	43*	48*	38*
	(-3.0)	(15.9)	(18.6)	(12.6)	(25.9)
$C^a - C_{H4M}^m$	-19*	30*	40*	39*	34*
	(-3.1)	(17.5)	(19.4)	(11.4)	(26.4)
$C^a - C_{H6M}^m$	-19*	32*	44*	48*	39*
	(-3.1)	(18.7)	(22.0)	(14.7)	(30.5)
$\lvert C^a - C_{LI}^m \rvert$	50	32	40	46	39
$\lvert C^a - C_{IWI}^m \rvert$	49	35	43	49	42
$\lvert C^a - C_{AWI}^m \rvert$	52	41	50	64	50
$\lvert C^a - C_A^m \rvert$	174	117	97	107	108
$\lvert C^a - C_{W7D}^m \rvert$	53	69	108	119	95
$\lvert C^a - C_{H7D}^m \rvert$	53	68	105	112	92
$\lvert C^a - C_{H2M}^m \rvert$	50	52	75	90	70
$\lvert C^a - C_{H4M}^m \rvert$	50	48	66	76	62
$\lvert C^a - C_{H6M}^m \rvert$	50	50	67	78	63
Percentage of cases in which C^a does not deviate from boundaries					
$C_I^l < C^a < C_I^h$	8.4%	14.6%	15.6%	9.0%	13.6%
$C_{AWI}^l < C^a < C_{AWI}^h$	6.8%	12.5%	12.2%	8.3%	11.3%
$C_{IWI}^l < C^a < C_{IWI}^h$	2.5%	13.4%	15.4%	10.9%	13.3%
Sample size	119	1,178	1,426	770	3,493

[a] See note for Table 3.

The model prices based on WSD7D and HSD7D seem to be the best predictors of the price of the BIO. However, because of trade in the TASE, better predictors were available for TASEO prices, namely, the AWISD, followed by the IWISD and LISD.[17] This ranking differs from Beckers' (1980) findings (for stock options), probably because of the low liquidity of the BIO.

Note that the average absolute deviation between actual price and model price decreases as the historical horizon increases, no matter what the depth and the maturity of the option. A similar result was found for the (relative) average deviation. However, the AWISD still provides the best prediction.

We now turn to test whether the deviations of the model price from the actual price is systematically related to any observable variables, especially the inputs of the B-S formula. If the minimal average deviation is systematically related to some variables, this fact can be utilized to minimize future biases and re-estimate the ISD. To this end, we first run the following regressions:

$$(1)\ (C_t^a - C_t^m) = \alpha_0 + \alpha_1\, C_t^m + \vartheta_t$$

$$(2)\ (C_t^a - C_t^m) = \alpha_0 + \alpha_1\, C_t^m + \alpha_2\, \tau_t + \alpha_3 \ln(S_t^r / K) + \alpha_4\, (\tau_t r_t^d) + \alpha_5\, (\tau_t r_t^f) + \alpha_7\, (\sqrt{\tau_t}\, \sigma_{W,t} / 2) + v_t$$

C_t^m is calculated by using the SD estimator that minimizes the average deviation of the model price from the actual price (see above). A significant positive coefficient for a parameter means that the model underprices options with high values of this parameter. Next, we test the following nested joint null hypotheses whereby the model provides unbiased estimates of the actual premium, otherwise this bias is systematically related to the inputs of the B-S formula:

$$(3)\ H_0^1 : \alpha_0 = \alpha_1 = 0$$

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[17] One could also use LISD for the BIO, but only two figures a week are available at best.

(4) $H_0^4 : \alpha_0 = \alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_7 = 0$

The results are summarized in Table 11.

We find that the constant term is always positive and significant, while the coefficient of the model price is always negative and significant. These findings can be explained by the fixed cost of transaction. An overpriced option does not necessarily introduce arbitrage opportunities owing to fixed transaction costs, especially when the price is low. But when prices increase, the fixed cost bar is eliminated, and consequently overpricing should be smaller. The coefficient of the model price is never insignificant, which means that the B-S model misprices the option.

The depth and maturity coefficients are significantly negative, implying that they have the right sign, as expected. Since the constant is significantly positive, as the option becomes ITM or draws close to maturity its price deviation from the model price decreases. Similar results are obtained for the WSD7D.

The coefficient of the options always decreases (increases in absolute value) when other parameters are added to the regression, while the constant increases. In other words, there is some rotation in the regression line. This might indicate errors in the variables, in which case instrumental variable techniques are required to perform the appropriate estimation. This observation can be found in most studies, but has never drawn much attention.

Table 11. Hypothesis Test: Model Price is Unbiased Estimator of Actual Price [a]

	TASEO		BIO					
	All maturities		3-months		6-months		12-months	
	$C^a - C^m_{AWI}$		$C^a - C^m_{H7D}$		$C^a - C^m_{H7D}$		$C^a - C^m_{H7D}$	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Constant	6.4*	29*	166*	184*	401*	203*	489*	1,459*
	(3.73)	(6.39)	(17.0)	(7.9)	(21.6)	(2.1)	(11.9)	(7.1)
C^m_t	-0.01*	-0.12*	-0.29*	-0.76*	-0.89*	-0.96*	-0.29*	-0.83*
	(-7.27)	(-15.1)	(-17.9)	(-21.2)	(-22.8)	(-26.7)	(-9.11)	(-7.8)
$\ln(S^r_t / K)$		-2,242*		-11,061*		-5,811*		13,906*
		(-13.4)		(-15.2)		(-3.5)		(6.7)
τr^d		2,740*		18,600*		8,156*		12,757*
		(6.2)		(17.7)		(4.6)		(6.2)
τr^f		1,114		-24,126*		-10,388*		-22,646*
		(0.9)		(-13.6)		(-3.1)		(-7.5)
$\sqrt{\tau}\sigma_A$		-180*		-178		2,558*		-2,911
		(-1.9)		(-0.6)		(6.9)		(-1.3)
$\sqrt{\tau}\sigma_{W7D}$		-204*		-1,388*		-295		-2,551*
		(-1.8)		(-4.2)		(-0.7)		(-3.1)
$\sqrt{\tau}\sigma_{H2M}$		-33		323		1,696*		97
		(-0.1)		(0.7)		(2.4)		(0.1)
τ		-169*						
		(-3.6)						
R^2	0.01	0.08	0.25	0.53	0.86	0.95	0.38	0.68
F	53	36	319	152	523	223	83	41
Sample size	3,463	3,463	943	943	83	83	137	137

[a] See note for Table 2 and Table 3.

(c) Dollar-dynamic neutral hedges

If the model price deviates from the actual price, the reason might be that the model accurately estimates the fair value whereas the market determines actual prices inefficiently. If so, above-normal profits might exist. The strategy would then be to construct a portfolio of options and maintain it at a given initial position until the expiration date. The dynamic riskless delta-neutral portfolio is achieved through daily changes in the dollar's position, thereby maintaining a short (long) $N(d_{1,t})$ dollars per unit of option (for one dollar underlying currency) bought (sold).[18]

The literature contains three strategies that differ in their portfolio structure and assumptions about the option price on the day of issue and thereafter (see Table 12),[19] according to the different hypotheses tested. The present study decomposes the total daily excess hedge return of the strategies into two components: (i) the daily excess return on hedging; and (ii) the deviation of the model price from the actual price on the purchasing day. The first component tests whether the options pricing model (OPM) can be used to maintain a neutral hedge; the second tests whether the OPM can detect over and underpriced options. If the OPM is valid, then both components should equal zero by the arbitrage argument. The strategies outlined in Table 12 differ only with regard to the second component: in the first strategy, this component (the deviation of the model price from the actual price on the purchasing day) is zero; the second strategy uses the real value of the deviation; the third strategy uses its absolute value. Analyzing this component is therefore meaningless in the first strategy (it is always zero), and the analysis with respect to the second and third strategies was performed in the previous section. All that remains is to analyze the first component — the daily excess return on hedging.

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[18] Once the position is put into effect, the dynamic neutral hedge can be achieved independently of the option. This is a crucial feature, since some options are not traded frequently.

[19] A fourth strategy was suggested by Galai (1978); it is not considered here since it is erroneous (see Haj-Yehia, 1997).

Table 12. Neutral Hedged Strategies Testing the Validity of the Option Pricing Model [a]

Strategy I	Portfolio of options	All options
	Option price on the issuing day	Model price
	Option price after the issuing day	Model price
	Excess hedge return on day t	$R_t = (C_t^m - C_{t-\Delta t}^m) - N(d_{1,t-\Delta t})*(S_t - S_{t-\Delta t})e^{r^f_{t-\Delta t}\Delta t} - (C_{t-\Delta t}^m - N(d_{1,t-\Delta t})*S_{t-\Delta t})(e^{r^d_{t-\Delta t}\Delta t} - 1)$
	This strategy tests whether	The B-S model fares well enough empirically, in spite of possible deviations from its assumptions
Strategy II	Portfolio of options	All options
	Option price on the issuing day	Model price
	Option price after the issuing day	Model price
	Excess hedge return on day $t_1 = t_0 + \Delta t$	$R_{t_1} = (C_{t_1}^m - C_{t_0}^a) - N(d_{1,t_0})*(S_{t_1} - S_{t_0})e^{r^f_{t_0}\Delta t} - (C_{t_0}^a - N(d_{1,t_0})*S_{t_0})(e^{r^d_{t_0}\Delta t} - 1)$
	Excess hedge return on day t (t>t₁)	$R_t = (C_t^m - C_{t-\Delta t}^m) - N(d_{1,t-\Delta t})*(S_t - S_{t-\Delta t})e^{r^f_{t-\Delta t}\Delta t} - (C_{t-\Delta t}^m - N(d_{1,t-\Delta t})*S_{t-\Delta t})(e^{r^d_{t-\Delta t}\Delta t} - 1)$
	This strategy tests whether	Writers receive, on average, too low or too high a price for their options
Strategy III	Portfolio of options	Long (short) options that are more than 5% underpriced $\left(0.95C_{t_0}^m > C_{t_0}^a\right)$ or overpriced $\left(1.05C_{t_0}^m < C_{t_0}^a\right)$
	Option price on the issuing day	Model price
	Option price after the issuing day	Model price
	Excess hedge return on day $t_1 = t_0 + \Delta t$	$R_{t_1} = (C_{t_1}^m - C_{t_0}^a) - N(d_{1,t_0})*(S_{t_1} - S_{t_0})e^{r^f_{t_0}\Delta t} - (C_{t_0}^a - N(d_{1,t_0})*S_{t_0})(e^{r^d_{t_0}\Delta t} - 1)$
	Excess hedge return on day t (t>t₁) [b]	$R_t = (C_t^m - C_{t-\Delta t}^m) - N(d_{1,t-\Delta t})*(S_t - S_{t-\Delta t})e^{r^f_{t-\Delta t}\Delta t} - (C_{t-\Delta t}^m - N(d_{1,t-\Delta t})*S_{t-\Delta t})(e^{r^d_{t-\Delta t}\Delta t} - 1)$
	This strategy tests whether	Opportunities exist for above-normal profits

[a] Δt is assumed to be a small time interval. It is also assumed that an option is exercised on the expiration date, at which time its value is max(0, S–k), where S is the exchange rate delivery settlement, and K is the strike exchange rate.

[b] By changing the sign we get the hedge return for overpriced options.

Decomposing profits into two components, as suggested above, is very useful, as the separate analysis of each component can be used to derive many more advanced strategies:

(a) Investors can decide on the magnitude of mispricing (which can also depend on depth and maturity) that should trigger initiating an arbitrage position on options.

(b) If an investor holds an inventory of options regardless of their price, he can decide whether to use the OPM to maintain a neutral hedge, which can depend on the depth and maturity too.

(c) Separation makes the results of testing the excess return on hedging more robust, since it does not depend on *any* assumptions about the options market.

(d) Separation enables utilizing panel data analysis.

(e) The decomposition overcomes the low liquidity problem and the nonsynchronization of the currency option market, because the daily excess returns on hedging do not use actual currency option prices.

We analyze only the daily excess return on hedging on date t for each listed option that we found to be mispriced, as follows:

$$(5)\ R_t = (C_t^m - C_{t-\Delta t}^m) - N(d_{1,t-\Delta t}) * (S_t - S_{t-\Delta t})e^{r_{t-\Delta t}^f \Delta t} - (C_{t-\Delta t}^m - N(d_{1,t-\Delta t}) * S_{t-\Delta t})(e^{r_{t-\Delta t}^d \Delta t} - 1)$$

for underpriced options; changing the sign gives the return on hedging for overpriced options.

This formula is further modified so as to bring the B-S model closer to reality. The simulated strategy requires only that the trader should adjust his inventory of dollars and bonds in such a manner that the interest rate earned on bonds from the previous day's trading continues to be fixed at the level of their day of purchase. Adjustment occurs if and only if the currency market, foreign bond market, and domestic bond market are all active simultaneously. The option position is put into effect when a mispricing is first detected in the course of a trading day in which all the above markets are active simultaneously.[51] The initial option position is maintained until the expiration date (we consider only options that expire during the sample period, because options must be held until expiration in order to examine the entire profit from an attempt to exploit mispricing).

This modification has at least four important advantages, besides the adjustment to the specific case of Israel.[52] First, it forces the strategy to comply with the B-S assumption of fixed interest rates. Secondly, this modified strategy is more realistic, especially the imposition of holding the option until the expiration date, because when an investor buys MAKAM (a short-term zero-coupon interest rate government bond) and holds it until expiration, his yield-to-maturity is predetermined on the day of purchase. Thirdly, according to this strategy the investor bears no interest-rate risk, which paves the way for considering models that employ non-stochastic interest rates. Finally, marginal adjustments minimize transaction costs.[53]

Table 13 shows average daily and total excess profits per contract (DEP and TEP, respectively) from utilizing mispricing by using ISD and HSD7D.[54] The above-normal average daily profits are negatively correlated with the depth (S/K) and maturity (τ) of

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[51] Therefore, all transactions are conducted only during Monday through Thursday.

[52] For example, bonds, options and currency markets in the world always have the same trading day, while in Israel Sunday is a trading day for NIS bonds and currency options, but not for foreign currency.

[53] If transaction costs are to be accurately considered, the Delta should also be adjusted. This will complicate the calculations even for very simple cases (see Merton, 1991, p. 380). However, since the transaction costs of thc most efficient sector (banking) are negligible, the results of the tests are valid.

[54] Very similar results are obtained when employing other standard deviation estimators. Rates of return cannot be reported, since the stratcgies are self-financed.

the option; they are positive for OTM BIO, and negative for ATM and ITM BIO, but always negative for the TASEO.[55]

Table 13. Above Normal Daily Profit from Dollar Dynamic Neutral Hedging[a]

	Total	Depth S/K				Time-to-maturity - τ (months)			
		OTM	ATM	ITM	Deep-ITM	0-0.25	0.25-2	2-4	4-6
3-month									
DEP_{I}	-0.67*	0.07	-0.40*	-1.25*		-0.94*	-1.17*	0.70*	1.87*
DEP_{H7D}	-0.60*	-0.16	-0.35*	-1.13*		-0.65*	-1.15*	0.66*	1.73
TEP_{I}	8.92*								
TEP_{H7D}	6.53*								
6-month									
DEP_{I}	-1.22*	0.25	-0.20*	-0.97*	-2.47*	-1.28*	-1.15*	1.42	
DEP_{H7D}	-1.16*	0.01	-0.25*	-1.00*	-2.22*	-1.16*	-1.38*	1.94	
TEP_{I}	13.44*								
TEP_{H7D}	-0.94								
12-month									
DEP_{I}	-7.47*	0.90	-9.14*	-3.18*	-10.04*	-5.08*	-7.26*	11.69*	13.00*
DEP_{H7D}	-7.35*	1.74	-0.94*	-3.21*	-9.91*	-4.78*	-7.39*	11.74*	12.73*
TEP_{I}	35.83*								
TEP_{H7D}	27.26*								
TASE									
DEP_{I}	-23.29*	-16.18	-10.94*	-24.73*	-49.24*	-17.30*	-29.15*	2.00*	
DEP_{H7D}	-23.14*	-15.20	-11.90*	-23.55*	-49.27*	-18.02*	-28.42*	1.57*	
TEP_{I}	6.39								
TEP_{H7D}	23.58								

[a] Average excess hedge return for each adjustment interval. DEP = Daily Excess Profit, TEP = Total Excess Profit. See also notes to Table 2 and Table 3.

Thus, investors were apparently more likely to exploit arbitrage opportunities from mispricing when options are more OTM and far from maturity. This finding, combined with the previous findings (that OTM far-from-maturity options were overpriced) lead

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[55] Note that the results for average daily profits are clearly evident even in the case of very small intervals of depth and maturity. Furthermore, their distribution is not far from normal and is not very skewed.

to the conclusion that investors were more likely to short option in order to exploit arbitrage opportunities, if they existed. But even if such opportunities did exist, investors may not have been able to exploit them because of strict restrictions on writing and short TASEO, low liquidity for far-from-maturity options, and the fact that the BIO could only be "long". Therefore, while the banking sector did enjoy arbitrage opportunities, other sectors had a far more limited scope of arbitrage opportunities, especially when transaction costs are taken into account. The behavior of average daily and total profits was the same under all the options. This means that it did not depend on the strike price, which was the only difference between them.

Although average daily excess profits are negative and similar using either ISD or HSD and ASD, total excess profits are positive, higher for ISD than for HSD, and increase when we refer only to above ±5% mispricing. These findings are similar to those reported by Chiras and Manaster (1978) and Latane and Rendleman (1976) for stock options. Combining this result with the previous result, that the AWISD is the best predictor of the option price, we conclude that the weighted ISD based on the B-S model is useful not only in determining properly hedged positions, but also in identifying relatively over- and under-priced options.

4. The Liquidity Premium

The 3- and 6-month BIO can be matched with TASEO on a given day of issue. These two options are identical in all respects but one — the former is non-tradable and the latter is tradable — so that the difference between their prices on the same day of issue represents the liquidity premium. This premium should be positive for every single option and hence in the aggregate as well.

Since a 100% match of strike prices is not very likely, I consider options to be matched if the best match between options' strike prices is within a range of ±5%. On the other hand, it is quite likely that the expiration dates of the matched options differ by up to three days, but these differences should not have any significant effect on the pricing. I ignore the 6-month BIO because only three matched observations were found, compared with 68 matched observations between the 3-month BIO and TASEO. The *average* liquidity premium is *positive*: NIS80 per one TASEO (or per 10 BIO). However, surprisingly, the liquidity premium was not positive for all options, which might indicate market inefficiency.

5. Summary and Conclusions

An examination of the boundary arbitrage and dominance conditions for rational options pricing leads us to reject the null hypothesis — that the currency options market is weakly efficient — only for the household sector and for the very near (under two weeks) maturity and deep in the money currency options.

We also test the strong-efficient market hypothesis, which is equivalent to testing the B-S model validity hypothesis given that the B-S model's assumptions are robust. We reject this hypotheses, since we find that (i) the B-S model underestimated ASD, (ii) actual prices were different from the B-S model prices, and (iii) the banking sector could have used the B-S model to utilize currency option mispricing in order to generate total positive profits by pursuing a neutral hedging strategy based on the B-S formula, especially in the case of far-from-maturity and OTM options. Another indicator of market inefficiency is the negative liquidity premium found for some BIO. However, it is worth noting that investors other than the banking sector had a limited scope of arbitrage opportunities, especially due to transaction costs, restrictions on short TASEO, and the low liquidity of far-from-maturity currency options.

The market underestimated the price of the TASEO relative to the model price based on the actual standard deviation. Investors were more influenced by the latest news and developments, and employed a standard deviation calculated as a weighted average of the historical standard deviation based on the most recent seven historical days and the expected standard deviation; they assigned less weight to the standard deviation based on the most recent historical seven days as the option approached maturity. We also find that the arithmetic weighted implied standard deviation based on the B-S model was useful, not only in determining properly hedged positions, but also in identifying relatively over- and under-priced options.

The BIO were either subsidized (in which case they did well) or over-valued (in which case they failed).

Our results resemble those found by Shastri and Tandon (1985) and Bodurtha and Courtadon (1986) for the Philadelphia Stock Exchange options market, where most of the deviations from the arbitrage boundary conditions were found in near-maturity currency options. Our results contradict the findings of Azolai and Landskroner (1993), who detected many deviations for all BIO types and sectors (banks and households). These differences may stem from changes in the average volatility of the exchange rate, better estimation and understanding by market participants than in its very early days, and the superior proxies used in the present study. Another reason may be that Azolai and Landskroner (1993) considered only ATM 3-month BIO, which is more vulnerable to deviations from arbitrage boundary conditions (see Haj-Yehia, 1997).

Appendix 1. Notations and Parameters

a. Notations[56]

Time

t= current time (date).

τ= **t**ime to maturity.

Options price

C^a = **a**ctual **c**all close option price.

C^h = **a**ctual **c**all **h**igh option price.

C^l = **a**ctual **c**all low option price.

C^m = **m**odel **c**all option price.

K= stri**k**e exchange rate.

Exchange rate

S^r = **r**epresentative **s**pot exchange rate.

S^a = **a**sk **s**pot exchange rate.

S^b = **b**id **s**pot exchange rate.

Interest rate

r_l^d = **d**omestic nominal interest **r**ate for **l**enders.

r_b^d = **d**omestic nominal interest **r**ate for **b**orrowers.

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[56] Bold letters are used as a mnemonic device.

r_l^f = foreign nominal interest rate for lenders.

r_b^f = foreign nominal interest rate for borrowers.

Option transaction cost

T_g^c = transaction cost for guarding a call.

T_e^c = transaction cost for exercising a call.

T_b^c = transaction cost for selling (biding) a call.

T_a^c = transaction cost for purchasing (asking) a call.

Foreign currency transaction cost

T_b^f = commission for selling (biding) one unit foreign currency.

T_a^f = commission for purchasing (asking) one unit foreign currency.

$T_s^f = S_a / S_r = S_r / S_b$ = proportional bid ask spread.[57]

Notes

1. The terms sale (bid) and purchase (ask) are from the viewpoint of the investor.
2. When a sub- or superscript is omitted the notation refers to either of the sub- or superscripts, or there is no difference between them.
3. The spot exchange rate is stated as one unit of foreign currency in terms of units of the domestic currency.
4. Interest rates and yields to maturity are for a period that matches the life of the option and are converted to a continuously compounded rate.
5. Transaction costs are stated as a percentage of the representative spot exchange rate.

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[57] The bid-ask spread is symmetric.

b. Parameters[58]

$N(.)$ = cumulative normal distribution.

$$d_1 = \frac{\ln(S/K) + [r^d - r^f + (\sigma^2/2)]\tau}{\sigma\sqrt{\tau}}.$$

$d_2 = d_1 - \sigma\sqrt{\tau}$.

$Z_{(1-\alpha)}$ = the inverse of the normal distribution with $1-\alpha$ cumulative distribution.

$\lambda = \frac{K_3 - K_1}{K_3 - K_2}$, where K_1, K_2 and K_3 are strike prices of three options priced C_1, C_2 and C_3, respectively.

σ = instantaneous standard deviation of the rate of return on the exchange rate.

$\sigma_{I,t} = f^{-1}(C_t^a, S_t, K, r_t^f, r_t^d, \tau_t)$ = **ISD** when substituting closing price of date t in the B-S formula.

$\sigma_{I,t}^h = f^{-1}(C_t^h, S_t, K, r_t^f, r_t^d, \tau_t)$ = **ISD** when substituting high price of date t in the B-S formula.

$\sigma_{I,t}^l = f^{-1}(C_t^l, S_t, K, r_t^f, r_t^d, \tau_t)$ = **ISD** when substituting low price of date t in the B-S formula.

$\sigma_{I,t}^{MBI} = f^{-1}(C_t^{MBI}, S_t, K, r_t^f, r_t^d, \tau_t)$ = **ISD** when substituting the Bank of Israel's minimum premium of date t in the B-S formula.

$\sigma_{I,t}^{LISD} = f^{-1}(C_t^{LISD}, S_t, K, r_t^f, r_t^d, \tau_t)$ = **LISD** when substituting closing price of the last trading date before t in the B-S formula.

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[58] All the SDs relate to the rate of return on the exchange rate.

$\sigma_{IWI,t} = \frac{1}{n}\sum_{i=t-6}^{t} \sigma_{IWI,i}$ = **IWISD** when using the closing price of an option that was traded on a certain day and that has the same classification in the text tables; the formula uses the ISD that was available for the options in the most recent 7 trading days; thus, n might be less than 7 if the option was not traded in all most recent 7 trading days.

$\sigma^{h}_{IWI,t} = \frac{1}{n}\sum_{i=t-6}^{t} \sigma^{h}_{IWI,i}$ = **IWISD** when using the **h**igh price of an option that was traded on a certain day and that has the same classification in the text tables; the formula uses the ISD that was available for the options in the most recent 7 trading days; thus, n might be less than 7 if the option was not traded in all most recent 7 trading days.

$\sigma^{l}_{IWI,t} = \frac{1}{n}\sum_{i=t-6}^{t} \sigma^{l}_{IWI,i}$ = **IWISD** when using the **l**ow price of an option that was traded on a certain day and that has the same classification in the text tables; the formula uses the ISD that was available for the options in the most recent 7 trading days; thus, n might be less than 7 if the option was not traded in all most recent 7 trading days..

$\sigma_{AWI,t} = \frac{1}{n}\sum_{i=1}^{n} \sigma_{AWI,i}$ = **AWISD** when using close price of all the n options that were traded at a certain day and that have the same classification in the table.

$\sigma^{h}_{AWI,t} = \frac{1}{n}\sum_{i=1}^{n} \sigma^{h}_{AWI,i}$ = **AWISD** when using **h**igh price of all the n options that were traded at a certain day and that have the same classification in the table.

$\sigma^{l}_{AWI,t} = \frac{1}{n}\sum_{i=1}^{n} \sigma^{l}_{AWI,i}$ = **AWISD** when using **l**ow price of all the n options that were traded at a certain day and that have the same classification in the table.

$\sigma_{A,t} = \frac{1}{\tau}\sum_{i=t}^{t+\tau}(R_i - \overline{R})$ = ex-post **ASD** applied over the time to maturity.[59]

$\sigma_{W,t} = \frac{1}{7}\sum_{i=t-3}^{t+3}(R_i - \overline{R})$ **WSD7D.**

$\sigma_{H,t} = \frac{1}{H}\sum_{i=t-1-H}^{t-1}(R_i - \overline{R})$ = **HSD,**[60] where:

$R_i = \ln(S_i / S_{i-1})$ is the daily rate of return on the exchange rate,[61]

$\overline{R} = \frac{1}{H}\sum_{i=t-1-H}^{t-1} R_i$ is the average daily rate of return on the exchange rate, and

H=7, 60, 120, 180 calendar days, is the historical period (H7D, H2M, H4M, H6M, respectively).

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[59] This is the maximum likelihood estimator asymptotically efficient in the class of Consistently Uniformly Asymptotically Normal (CAUN).

[60] See footnote 59.

[61] Weekends and holidays can either be taken into account or ignored, and each approach makes its own implicit assumptions (see Campbell, Lo and McKinlay, 1997). As the volatility of weekends' and holidays' returns is not significantly higher than that of weekday returns, I ignore weekends and holidays. Galai (1978), for example, uses both.

Appendix 2. The Data[62]

The database contains five source files:[63]

1. TASEO

A daily database for the period April 10, 1994, to May 2, 1996, containing the following information for each option traded: Date of trade, strike date, strike price, opening price, closing price, type of closing price, high price, low price, number of transactions, number of contracts traded, trade volume in NIS, and number of contracts as open positions. I examine only call options, because there was very low activity in put options. There are 3,493 observations, each reporting the results of an option series traded on a particular date.

2. BIO[64]

All BIO auctions for the period January 2, 1991, to May 2, 1996:[65]

i. ATM 3-month BIO — January 2, 1991, to May 2, 1996; total: 691 auctions.

ii. Pure ATM 3-month BIO — December 1, 1992, to May 2, 1996; total: 332 auctions.

iii. Pure ATM 6-month BIO — September 8, 1994, to April 30, 1996; total: 87 auctions.

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[62] Irregularly sampling the data poses no problems for continuous time processes (see Campbell, Lo and MacKinlay, 1997, p. 363).

[63] The first file was obtained from the TASE; the rest — from the Bank of Israel.

[64] Excluding options with a low trading volume (in terms of number of contracts or dollars) is possible, but might create two problems: it would reduce the data sample and introduce data snooping. I therefore prefer to use all available data observations.

[65] The auctions held between November 1989 and December 1990 are not well documented.

iv. Pure ATM 12-month BIO — January 22, 1991, to August 30, 1994; total: 126 auctions.

For each auction the file contains:

i. Issuing and striking date.

ii. Duration.

iii. Strike exchange rate.

iv. Supplied, demanded and traded options.

v. Maximum, minimum, close and average premium.

vi. The Bank of Israel's minimal premium allowed.

3. Interest rates

Close annual quoted discount yield to maturity (YTM) for the period January 1, 1991, to June 1, 1996, as follows:[66]

i. MAKAM: Daily YTM for almost every period under 12 months.

ii. PAKAM: Weekly YTM for 1 week, 2 weeks, and 1 month.

iii. AZAQ: Monthly average.

iv. PAZAK: Monthly average.

v. LIBOR: Monthly average YTM for 1 day, 7 days, 1 month, 3 months, 6 months and 12 months.

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[66] See Glossary (p. 39 below) for an explanation about these types of securities.

The interest rates of different maturities are obtained by simple average interpolation based on the two flanking rates.

Table A1 provides the details about the way the interest rates were calculated.

Table A1. Types of interest rate

Type			Bank	Household
Domestic	Lending	r_l^d	MAKAM	PAZAQ
	Borrowing	r_b^d	MAKAM	AZAK
Foreign	Lending	r_l^f	LIBOR-0.125%	LIBOR-2%
	Borrowing	r_b^f	LIBOR	LIBOR + 3%

4. Exchange rates

Daily representative exchange rate for the period January 1, 1991, to June 1, 1996.[67]

5. Transaction costs

Transaction costs were stable throughout the period considered in this study. They differ between sectors and types of transaction, as shown in Table A2.

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[67] For a description of the Israeli exchange rate policy and regime, see Haj-Yehia (1992).

Table A2. Transaction costs [a]

Security	Investor	Ask	Bid [b]	Redemption [c]	Safekeeping [d]
MAKAM	Banks	0.0028[e, g]	0	0	0
	Institutes	0.05[c]	0	0	0.125
	Households	0.1[c]	0	0	0.125
		T_a	T_b	T_e	T_g
BIO	Banks	0	0	0	0
	Institutes	0.1[g]	0	0.1	0.1
	Households	0.25[g]	0	0.25	0.1
TASEO	Banks	1.775 NIS/op	1.755 NIS/op	0	0
	Institutes	1.5[b]	1.5	0	0
	Households	4[b]	4	0.25	0
Currency exchange fee		T_a^f	T_b^f		
	Banks	0	0		
	Institutes	0.1[b]	0.1		
	Households	0.15[b]	0.15		
Currency exchange spread		T_s^f	T_s^f		
	Banks	0	0		
	Institutes	0.25[b]	0.25		
	Households	1.25[b]	1.25		

[a] Percentages unless otherwise indicated; assuming that the minimal fee is never effective; and ignoring non-tangible costs in terms of opportunity costs.

[b] Fee is a percentage of the transaction value on the transaction day.

[c] Fee is a percentage of the redemption value on the transaction day.

[d] Fee is a quarterly percent calculated from the value on the date of expiration.

[e] Fee calculated in proportion to time to maturity. For example, if the MAKAM expires within 36 days, the fee for a household is: 0.1% × 365/36 = 1.014%.

[f] Changed to 0.002% for 1993, and to 0.001 for 1994.

[g] Fee is a percent of the exchange rate.

Source: Unpublished information obtained by the author from commercial banks and the Bank of Israel.

Appendix 3. Glossary

ASD	Actual standard deviation
ATM	At the money
AWISD	Arithmetically weighted ISD
AZAQ	*Ashrai Zman Katsar* [short term bank loan]
BIO	Bank of Israel currency options
B-S	Black and Scholes
CBOE	Chicago Board Option Exchange
DEP	Daily excess profit
ERDS	Exchange rate delivery settlement
HSD	Historical data standard deviation
HSD7D	Historical data standard deviation derived from most recent 7 days
HSD[i]M	Historical data standard deviation derived from most recent i months
ISD	Implied standard deviation
ITM	In the money
IWISD	Inter-temporal WISD
LIBOR	London Inter-Bank Offered Rate
MAKAM	*Milve Ktsar Mo`ed* [short-term zero coupon government bond]
NIS	New Israeli Shekel
OPM	Options pricing model
OTC	Over-the-counter
OTM	Out of the money
PAKAM	*Pikadon Ktsar Mo`ed* [short-term bank deposit]
PAZAK	*Pikadon Zman Katsuv* [fixed-term bank deposit]
SD	Standard deviation
TASE	Tel Aviv Stock Exchange
TASEO	Tel Aviv Stock Exchange currency options
TEP	Total excess profit
WISD	Weighted average ISD
WSD	Window standard deviation
YTM	Yield to maturity

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Chapter 3: The Robustness of Exchange Rate Models, Unit roots and Cointegration

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Abstract

Three basic models have been proposed to explain the exchange rate: Purchasing Power Parity (PPP), the Balassa-Samuelson model and the random walk model. The robustness of these models is not merely a statistical curiosity but has important implications in many economic and financial models. During the last three decades this issue has attracted many researchers who have used the latest econometric methods. However, there are many who are skeptical, since most of the findings rest on the razor's edge. The purpose of this paper is threefold: first, to summarize the empirical literature on modeling the exchange rate process; second, to suggest an appropriate transformation of the aforementioned models into an econometric specification; and finally to introduce a precise framework for examining their separate or "pooling" validity. In particular, I use a formal Complete Vector Error Correction Model (CVECM) along with other statistical methods. Applying this methodology to the 1920s float of US Dollar-Britain Sterling exchange rate, I find evidence that cannot reject the hypothesis of the invalidity of the long run PPP with short run deviations.

JEL classification: F31, F41.

Keywords: Purchasing Power Parity (PPP), the Balassa-Samuelson model, random walk model, and VEC Model.

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1. Introduction[*]

The history of exchange rate modeling can be partitioned into two periods. The first period ends in mid 1970s. During this period, debated issues were based on the *theoretical validity* of the models, ranging from which prices are appropriate in asserting the theory to whether the theory itself is a truism (see reviews by Frenkel, 1976, Genberg, 1978, Kalamotousakis, 1978, and Officer, 1976). For the last three decades the debate has been focused on *empirical validity*. This paper will examine the robustness of the models in light of recent developments in the econometrics theory of time series analysis.

I shall show in this paper that there are three principal hypothesized processes for the exchange rate. The first is the Purchasing Power Parity (PPP) theory which holds that the real exchange rate is time-stationary or in econometric terms that its unconditional expectation is constant over time and its projection converges to this value. The second hypothesis assumes that the exchange rate exhibits a (deterministic) trend-nonstationary process, such that if one subtracts the time trend, the result is time-stationary.[1] The third theory assumes that the exchange rate process is (stochastic) drift-nonstationary, such that only first (or higher) differences of the exchange rate are time-stationary.[2]

The determination of the true process of the exchange rate is important for both theoretical and practical reasons. On the theoretical level, the PPP theory is a major building block of a plethora of theoretical and empirical models of the exchange rate (see Dornbusch, 1976, essays in Frenkel and Johnson, 1978, and Mussa, 1982). At the same time, theories other than the PPP, especially those that assume a random walk

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[*] I thank Olivier Blanchard, Victor Chernozhukov, Jerry Hausman, Guido Kuersteiner, and Whitney Newey for their helpful advice and comments. I would also like to thank the participants in the econometrics lunch at MIT. All remaining errors are mine.

[1] By stationarity I mean variance-stationarity.

[2] The terms stochastic drift nonstationary, unit root, and integrated process have the same meaning, and will be used interchangeably.

process, are also commonly used as a basic assumption in many theoretical and empirical models of financial assets. Findings in favor of one theory will cast doubts on the other and undermine the theoretical and empirical models based on that theory.

On the practical level, evidence in favor of one of the models will mandate its serious consideration in macro-economic policy issues, such as whether there is a case for managed float and whether a floating exchange rate regime neutralizes foreign shocks (see Durlauf, 1989, and Frenkel, 1981a). Further, it is essential to know the underlying model when considering dynamic multipliers and when forecasting the exchange rate or its errors (see Hamilton, 1994, Ch. 15). Furthermore, the true process of the exchange rate is a crucial issue in investment decisions and hedge strategies (see Hull, 2002, and references cited there).

In addition, implications of the exchange rate analysis can be applied in the analysis of other common stylized facts in macroeconomic time series behavior (see Enders, 1995, Ch. 3). In our instance, evidence in favor of the random walk is not peculiar to the exchange rate. Nelson and Plosser (1982) have found evidence that GNP and other macroeconomic time series behave like random walks. Since then, the literature in time series analysis is voluminous and enormously complex. Lessons from investigating the exchange rate and price indices processes and their relationship can assist in the investigation of other macroeconomic time series.[3]

There are at least three major contributions of this paper. First, I provide a statistical and econometric methodology that, I argue, should be employed when analyzing time series. In particular, this paper provides a natural framework that fits exactly into the analysis of the exchange rate models by employing a formal Complete Vector Error Correction Model (CVECM) analysis and several competing statistics for precise inference. The CVECM and some of these statistics and procedures are employed for the first time in the analysis of the PPP and random walk theories.

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[3] The bibliography in Judge et al. (1985) on this topic alone runs to more than 13 pages even though it is an early book.

Second, the various theoretical models are translated into a nested econometric model which makes them testable for robustness and exposes them to systematic empirical evaluation using the proposed CVECM analysis. This approach provides direct inference and comparison of the three models within the same framework and enables the examination of their simultaneous validity.

Finally, I will present a survey of the empirical research for the last two decades. This research has focused on the adequacy of the econometric techniques used for testing the validity of the monetary approach to the PPP. The most important points are highlighted. In addition, I proceed to apply the above methods to the 1920s' experience of floating USD-GBP exchange rate.

This paper is organized as follows. Section 2 outlines the theories for modeling the exchange rate process and provides an econometric interpretation for the economic theory. Section 3 introduces the debate in the literature concerning the empirical methodology and statistical inference used in testing these models. Section 4 offers an econometric methodology that nests all the proposed models from the exchange rate and time series analysis literature. An application of this methodology for the 1920s floating regime is then discussed in section 5. Concluding remarks are presented in section 6. In general, imprecise or inappropriate parameter estimates have prevented a consensus being reached on the empirical findings. This is mainly due to the fact that in testing the validity of the PPP and competing theories, researchers have been limited by the selection of econometric techniques and statistics and the estimated regressions. This paper proposes a CVECM analysis as a natural framework for testing the robustness of the exchange rate models. Applying this method to the period 1921-1925 using monthly observations for the USD and GBP, I find evidence that *cannot reject* the hypothesis of *no* long run PPP model with short run deviations. However, this does not mean a rejection of the PPP.

2. The basic theory

This section introduces the main three theories for modeling the exchange rate: (i) the PPP model (ii) the Balassa-Samuelson model, and (iii) the random walk model. Econometric interpretation and transformation is presented in order to enable the empirical testing of their robustness.

In its simplest version, i.e. assuming perfect competition, identical preferences, no barriers to trade and no transaction costs, the long run PPP theory has two versions. The first is the *absolute* version which states that the long run real exchange rate is time invariant and equals its unconditional expectation. Simply stated, the price of a commodity is the same among countries regardless of the unit of measure ("Law of One Price"). This can be expressed as follows:

$$\textbf{(1)}\ e \equiv \ln(SP^{*} / P) = 0,$$

or equivalently:

$$\textbf{(2)}\ s = p - p^{*};$$

where uppercase letters express level and lowercase letters express log, e and s are the real and nominal spot exchange rates defined in local currency units per one foreign currency unit, and p and p* are the local and foreign price levels, respectively. Allowing for transaction costs, eq. (1) can be rewritten as:

$$\textbf{(3)}\ \ln(CSP^{*} / P) = 0,$$

or equivalently:

$$\textbf{(4)}\ s = c + p - p^{*};$$

where C and c are transaction costs in level and log, respectively.[4]

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[4] Eq. (3) implies the *exclusiveness condition*: in the long run, as long as PPP holds, the exchange rate should only be a function of relative prices.

The second version of the PPP is the *relative* version.[5] It asserts that equilibrium changes in the real exchange rate equal zero or, equivalently, that changes in the nominal exchange rate equal the changes in the difference between foreign and domestic prices. This can be expressed as follows:

$$\textbf{(5)}\ \Delta e_t = 0,$$

or equivalently:

$$\textbf{(6)}\ \Delta s + \Delta p^* - \Delta p = 0.$$

In both versions, the result is ensured by perfect commodity arbitrage (or equilibrium in the assets market) that acts as an error correction mechanism whenever a deviation occurs. Both versions are more robust in a hyperinflationary economy. When examining the exchange rate between countries with differing inflation rates, the relative version is likely to hold even if the absolute version does not hold at the same time.

The PPP doctrine was first proposed by Wheatley and Ricardo during the first part of the 19th century and by Cassel during the 1920s (see Dornbusch, 1987, and Frenkel, 1976, 1981a), though it was first mentioned by Malynes as early as 1601 (see Kalamotousakis, 1978). They have viewed the doctrine as an extension of the quantity theory of money in an open economy (the monetary approach) and believed it to be valid for both the short and long run (see Dornbusch, 1987, Frenkel, 1976, Genberg, 1978, and Officer, 1976). By now, the basic consensus seems to be that the PPP theory is supposed to hold in the *long run* (though possibly with long and variable lags) as long as the shocks are of a *monetary* as opposed to *real* origin (see Daniel, 1986b, Davuytan and Pigginger, 1985, Dornbusch, 1980, 1985, Frenkel, 1976, 1978, Gailot, 1970,

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[5] "Indeed, much of the post WW1 debate over re-establishing pre-war parities, which provided the genesis of PPP theory, implicitly referred to relative PPP," Froot and Rogoff (1994), p. 5.

Hodgson and Phelps, 1975 and Miller, 1986). Real shocks affect equilibrium relative price structure and bring about price adjustment.[6]

PPP is not expected to hold in the short run due to slower adjustment in commodities markets as compared to assets market (see Adler and Lehmann, 1983, Artus, 1978, Choudhry et al., 1991, Dornbusch, 1980, Frenkel, 1981a, Kravis and Lipsey, 1978, Mussa, 1982, and Taylor and McMahon, 1988). This occurs principally during periods that are dominated by ample news and by government intervention that alter expectations and cause high volatility in the exchange rate. This is due to the higher sensitivity of exchange rate to expectations relative to prices. Another theoretical reason for deviation from PPP in the short run has been expressed by Daniel (1986b) and Dornbusch (1980), who claimed that if commodity prices are sticky, then changes in the nominal exchange rate will induce changes in the real exchange rate causing deviation from PPP (see also Davuytan and Pigginger, 1985).

A different hypothesis regarding the PPP assumes that the exchange rate behaves as a nonstationary process, such that the PPP is likely to be violated even in the long run. This viewpoint has been expressed in the seminal papers of Balassa (1964) and Samuelson (1964). They argue that cross-country differential sectoral rates of productivity growth (especially between traded and non-traded goods) change real costs and relative prices and therefore bring about divergent movements in exchange rate adjusted national price levels. Therefore, the relative price level is high in wealthy countries as compared to poor countries and rises rapidly in rapidly growing countries relative to slow growing ones (see Officer and MacKinnon, 1971). This will cause the exchange rate to be deterministic trend nonstationary. If differential productivity shocks are permanent (or repeated), sectoral productivity shocks can induce a deterministic trend (or stochastic drift) in the real exchange rate.

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[6] For this to hold, some authors have argued that imperfect substitution between domestic and foreign commodities is required (see Flood, 1981, and Mussa, 1982). However, this seems to be a redundant assumption. Consider, for instance, changes in transport costs or tariffs (for empirical evidence, see Frenkel, 1981a).

Another version of this competing hypothesis assumes that the nonstationarity stems from a stochastic drift component, with the random walk process as its prototypical formulation.[7] The common underlying theoretical model is the Martingale model of exchange rate as an implication of financial arbitrage in the bond markets (see Adler and Lehmann, 1983, Apte *et al.*, 1994, Darby, 1980, and Hakkio, 1984).[8] A random walk process of exchange rate is a widespread assumption in pricing financial assets; especially currency options (see Biger and Hull, 1983, Black and Scholes, 1973, Cox and Rubinstein, 1985, and Garman and Kohlhagen, 1983).

There is a substantial theoretical and empirical literature that discusses additional fundamental factors that affect long run exchange rate equilibrium. This literature discusses supply factors, e.g. wealth redistribution, productivity, and strategic pricing decisions by firms and demand factors, such as government spending, income variables, and strategic consumption decisions by consumers. Froot and Rogoff (1994) present an excellent survey. Another stream of theoretical literature emphasizes Dornbusch-Krugman model of 'pricing to market' theory (see Dornbusch, 1987, and Krugman, 1987, Engle, 1993, Govannini, 1988, and Isard, 1977).

3. The empirical research[9]

The recent extensive empirical research during the last two decades on the exchange rate process has been concentrated mainly on the PPP theory and random walk model with little attention given to the models of Balassa (1964) and Samuelson (1964), though it has been recognized that changes in relative prices between tradable and

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[7] The random walk hypothesis has the disturbing implication that shocks to real exchange rate have no tendency to reverse, which clearly implies there is no tendency for PPP to hold.

[8] A martingale process of exchange rate is a stochastic process in which successive increments are unpredictable. In such a case, the martingale theory predicts oscillatory behavior around the PPP, with serially independent increments from the PPP (see Chow and Teicher, 1978, and Feller, 1966). This suggests that exchange rates follow a random walk process (for empirical results see Adler and Lehmann, 1983, Darby, 1980, and Frenkel, 1981a).

[9] See summary table in Appendix 4.

non-tradable goods will introduce errors into PPP.[10] The development of new statistical methods that have a bearing on the time series properties of the exchange rate and price indices as well as new databases that permit the examination of larger and more desegregated time series has attracted many empirical researchers to reassess the PPP vs. random walk although no consensus has been reached.[11] This has only intensified the claims for and against the validity of the PPP. Since the basic assumption of the theory is a free floating exchange rate, two primary periods of interest have been examined: the 1920s and the 1970s.[12] Nevertheless, some researchers have used sample periods incorporating both fixed and flexible exchange rate regimes, arguing that the theory holds during both (see Abuof and Jorion, 1990, Apte et al., 1994, Ardeni and Lothian, 1991, Froot et al., 1995, and Lothian and Taylor, 1994).

The debate among the researchers can be partitioned into two areas: I) attributing the evidence to *imprecise* parameter estimates; and II) attributing the evidence to *inappropriate* parameter estimates.[13] The former debate revolves around the appropriate econometric techniques and database that should be employed; the latter is concerned with the appropriate macroeconomic model that should be estimated. The following subsections discuss the evolution of this debate while focusing on methodology:

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[10] It seems that this is due to the difficulties in econometric identification of these changes due to their systematic relation with the exchange rate. A decrease in the price of non-tradable relative to tradable goods tends to be associated with depreciation. The effect of this change on chosen price indices proxy depends on the weights of these price indices. I think the GIVE method might help resolve this problem. For some measurement of the effect of non-traded goods prices on the real exchange rate see Engel (1995).

[11] Some researchers have obtained quite unambiguous findings to confirm and others to refute the PPP, and some researchers have even obtained contradicting evidence (within the same research).

[12] Following WWI and again in 1973, most countries abandoned the regime of pegged exchange rate (the Gold Standard and Bretton Woods, respectively) and switched to a regime of flexible (though managed) rates. In the late 1980s, countries began adopting a band regime (e.g. EMS, Snake and Louver Accord). It is worthwhile noting that Bretton Woods system was actually a band regime of ±1 percent around the dollar exchange rate.

[13] Since the 1920s, the econometric investigators have been disenchanted with time series models and their problems, e.g. Yule (1926).

conventional static tests, unit root (UR), cointegration (CI), Error Correction Model (ECM), Fractional Integration (FI), and Variance Ratio (VR). Empirical findings are of less concern, especially since most of them are proved to be a result of imprecise estimation.[14]

3.1. Imprecise parameter estimates

One standard view of the first school states that the PPP has fared well (or poorly) due to incorrect econometric estimation and the low power of the tests rather than positive (or negative) results for (or against) it. Hence, the researchers have provided suggestions for overcoming these problems.

Krugman (1978) and Frenkel (1981) have suggested the use of IV to overcome the bias in the coefficient of the bivariate estimation (see below). Hakkio (1984) has argued that in a multi-exchange rate world with highly integrated capital market, a multi-lateral exchange rate model must be preferable to a bilateral model. This is because a multilateral model increases the cross-sectional variability in the data and accounts for the large correlation between exchange rate movements. Similarly, Abuaf and Jorion (1990) have suggested pooling the data in a system of univariate autoregressions estimated jointly by the SURE method, in order to fully exploit the information in cross-equation correlation and also advocated the use of the statistics proposed by Dickey and Fuller (1979).[15]

Corbae and Ouliaris (1988) have argued that the nonstationarity in the level of the spot exchange rates and domestic and foreign price indices makes the use of "conventional" tests of the absolute version of PPP inappropriate. Therefore, the authors have suggested testing the absolute version via a cointegration system.

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[14] However, a summary of the empirical findings, methodology, and data sets is given in Appendix.

[15] However, the empirical distribution of the statistics has to be derived by simulation analysis as in Fuller (1976).

By comparing ECM technique with the conventional technique, Baillie and Selover (1987) have found that differences in the degree of integrability and lack of cointegration between the variables have resulted in the invalidity of the estimated models and therefore, their failure in forecasting. The authors have suggested applying ECM rather than the conventional techniques.

Ardeni and Lubian (1989, 1991) and Glen (1992) have suggested using the cointegration method and the VR of higher order difference. These methods do not rely on causality or exogeneity among the series and thereby avoid the potential bias.

Apte et al. (1994) has attributed the dismal performance of the standard OLS and GLS estimation to the lead and lag effects in non-traded vs. traded goods inflation. The authors have therefore suggested the capital market estimator as an instrumental variable which is tailored to this well-specified type of errors-in-variables. Further, they have constrained the SURE estimation procedure in an attempt to exploit a priori knowledge about the implications of the US dollar-based equations on the PPP relation between cross rates.

Fisher and Park (1990) have suggested employing the Park P-test (see below). Some have argued that this test has been stood on its head. Tests should impose symmetry and proportionality hypotheses rather than estimating them, and test for the stationarity of the real exchange rate rather than imposing it (see Adler and Lehmann, 1983, Darby, 1983, Edison, 1987, Frenkel, 1980, Hakkio, 1984, Meese And Rogoff, 1988). However, others have shown the low power of this test. Further, there exists a trade off between efficiency and consistency when imposing (non) stationarity. If the restriction is true, then the estimators, forecasts, and hypothesis testing are more accurate; otherwise, they will be unreliable no matter how large the sample.

Another proposition of the first school argues that if PPP is to hold only in the long run, then a meaningful test must meet some essential conditions. In the first place, the research must consider fairly long sample periods, allowing commodity prices to fully adjust and cover a number of cycles of deviations from PPP (which might take a number of years).[16] In the second place, researchers must employ fairly frequent observations (daily or weekly) in order to capture the low frequency component of the time series. For well known practical reasons these conditions have not been met in research (see Abuaf and Jorion, 1990, Ardeni and Lubian, 1989, 1991, Choudhry *et al.*, 1991, Edison, 1987, Edison and Klovland, 1987, Enders, 1989, Frankel, 86, Glen, 1992, Froot et al., 1995, Lothian, 1990, Kim, 1990, and Papell, 1994).[17] In addition to these papers, it is worth mentioning Froot et al. (1995) which employed a desegregated data set for a period of 700 years.

First school has also emphasized the well known index number problem. As a macroeconomic theory, the PPP uses a macro-aggregate price index. However, for this to be valid, the price index must use the same weights in both countries (see Balassa, 1964, for numeric evidence) and the "Law of One Price" holds exactly for every commodity in the price index though it is possible that deviations roughly cancel out when averaged (see Balassa, 1964, Davuytan and Pigginger, 1985, Engle, 1993, Frenkel, 1978, Froot et al., 1995, and Samuelson, 1964). Nevertheless, a wholesale price index (WPI) and/or a cost-of-living price index (CPI), though they are only crude approximations of a proper index and are weighted differently across countries, are commonly used as proxies in research.[18] For example, since WPI is biased towards traded and manufactured goods, tests based on CPI tend to reject less frequently than tests based on WPI (see Frenkel, 1978, Edison, 1985, and Taylor and McMahon, 1988). Edison (1987) is the only paper to have used a GDP price deflator. However, it is noteworthy that when money is neutral and price index movements are governed by

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[16] See Appendix 4.

[17] Froot and Rogoff (1994), subsection 2.3.3, have spelled out the importance of the duration of the data using numerical examples, though they associated it (mistakenly) with the power of the test.

[18] Starting in the 1950s, there have been some attempts to construct international price indices for identical baskets of goods (see Summers and Heston, 1991).

monetary shocks, then differential weighting does not matter and relative PPP should still hold.

3.2. Inappropriate parameter estimates

This school has also proposed the use of modified or alternative versions of the PPP in lieu of the native version.

Kravis and Lipsey (1978) and Genberg (1978) have given support for the view that emphasizes the intermediate run, as opposed to the Balassa-Samuelson secular hypothesis, and argued that exchange rate movements exert real effects because movements in prices and exchange rates are not synchronized.

Ardeni and Lubian (1989, 1991) and Edison (1985) have argued that by estimating the restricted equation arising from the theory without testing the imposed restriction, one risks using a mis-specified equation to assess the theory.

Edison (1987) has argued that the PPP doctrine must be enlarged to account for the effects of changes in structural factors. Therefore, a naive version of the PPP relationship does not adequately model the $/£ exchange rate.

Another point made by this second school is that there are factors that might explain deviation from the PPP but that do not imply its collapse. In particular, it has been found that the Canadian announcement in June 1961 of the intention to push the Canadian dollar to a significant but unspecified discount vice-à-visa the US dollar (Choudhry *et al.*, 1991), and Britain's announced intention to re-establish pre-World War I parity (Taylor and McMahon, 1988) prior to the official abandonment of the flexible exchange rate regime in May 1962 and June 1925, caused a deviation of the CAN$/US$ and £/US$ exchange rates from PPP. Frenkel (1981b) has demonstrated that such announcements caused the collapse of the PPP in the 1970s. On the other hand, Haj-Yehia (1992) has suggested that government intervention in Israel has helped maintain the PPP. Further, the Chilean (and more recently Brazilian) horizontal exchange rate band is intended to maintain PPP even in the short run.

Davuytan and Pigginger (1985) have argued that Frenkel‘s (1981a) observation that the PPP performed better during the 1920s than the 1970s is due to the increase in the volatility of real shocks rather than the collapse of the PPP.

Further we must take into account that the strict PPP is likely to perform better between neighbors as a result of low barriers to trade and a high level of competition. Houthakker (1978) has argued that transportation costs and other barriers to trade can cause persistent departures from PPP. In particular, Taylor (1988) has shown that transportation costs result in such a departure, which is implied by $\alpha_0 \neq 0$ as mentioned above. In this case, it is still of interest to test for cointegration of the exchange rate and price indices.

4. The econometric methodology

In attempting to test the validity of PPP, researchers have tested either the absolute or the relative version. The absolute version has been tested with the following trivariate regression:

$$\textbf{(7)}\ s_t = c_0 + \alpha_1 p_t - \alpha_2 p_t^* + \upsilon_t\,;$$

while the relative version has been tested by:

$$\textbf{(8)}\ \Delta s_t = c_1 + \alpha_1 \Delta p_t - \alpha_2 \Delta p_t^* + u_t\,;$$

with the null hypothesis $\alpha_1 = \alpha_2 = 1$ in both and also $c_1 = 0$ in the second case.[19] We do not require $c_0 = 0$ (and most likely it is not equal to zero) since this results from the existence of transaction costs and tariffs.[20]

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[19] The null hypothesis $\alpha_1 = \alpha_2$ is known as the "symmetry" hypothesis while the $\alpha_1 = \alpha_2 = 1$ is known as the "proportionality" hypothesis.

[20] Many research papers have incorrectly imposed $c_0 = 0$ and mistakenly refuted PPP when $c_0 \neq 0$.

The null random walk of the real exchange rate has been tested by applying the ADF-test, PP-test, VR or FI tests. Some of the very early efforts did not use modern unit root test methodology but did, nevertheless, illustrate the subtleties (see Adler and Lehmann, 1983, and Darby, 1983, among others).

The Balassa-Samuelson model has been examined by testing the significance of the fundamental factor in the deviation from PPP. Balassa (1964), for example, has regressed the following regression for a cross of twelve industrial countries for the year 1960:

$$(9) \left({}^{P}\!/_{sp^*} \right)_i = \mu + \eta \left({}^{GNP}\!/_{POP} \right)_i + \varsigma_i ;$$

where POP is population, ς is an error term and the subscript i denotes country.

I will argue that the above approaches are incorrect and therefore imprecise estimation may be leading to incorrect inferences.

Researchers are not free to select the version, but rather the version should be dictated by the stationarity and cointegration of the variables. On the one hand, if variables are nonstationary and cointegration exists, deviation from the long run equilibrium ought to be significant and have explanatory power in eq. (8). Further, if cointegration exists, it must be that the innovations are non-invertible and consequently the standard VAR representation cannot be recovered from the MA as in eq. (8). Therefore, omission of the long run deviation will cause misspecification (see Hamilton, 1994, Ch. 19.1). On the other hand, if variables are nonstationary and cointegration does not exist, then eq. (7) and (9) will yield spurious results and cause false inferences; in addition, imposing a theoretical restriction on the cointegrating vector will be imprecise (see Granger and Newbold, 1974). Furthermore, if variables are stationary, then eq. (8) is over-differenced which will result in inconsistent estimators (see Fuller, 1995), whilst differencing a trend stationary series or de-trending a drift stationary series are inappropriate procedure.

The properties of the statistics and therefore the statistical inference are also dictated and affected by the stationarity and cointegration of the variables. Nonstationarity in the level of the exchange rates and price indices makes use of conventional statistics for the absolute version of PPP inappropriate and misleading. For instance, nonstationary variables or variables exhibiting a near random walk are likely to display a high R^2 with highly correlated residuals even though little (or no) correlation exists between the variables.[21] However, researchers have not paid sufficient attention to this warning of spurious regression, nor to its effect on the asymptotic properties and distributions of the statistics which will lead to imprecise inference.[22] Including regressors that are not in the true data generating process (DGP) or excluding regressors that present in the true DGP decrease the power of the unit root tests, specially when the omitted variable is trended and the sample size increases or the omitted variable is significance (see Campbell and Perron, 1991). "The key problem is that *the tests for unit roots are conditional on the presence of the deterministic regressors and tests for the presence of the deterministic regressors are conditional on the presence of a unit root*," (italics are in origin) as stated by Enders, p. 255.[23]

Most of the findings regarding the PPP or random walk rest on the razor's edge. The statistical tests for detecting the kind of stationarity are not sufficiently sensitive and have a low power, especially in the case of a near-unit root and board line cases (see Hamilton, 1994, p. 515, Phillips and Perron, 1988, and Schwert, 1989). Stationarity, trend stationarity, and drift stationarity differ in their implication at infinite time

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[21] This phenomenon of spurious regression was first discovered in the seminal paper by Granger and Newbold (1974) and later explained theoretically by Philips (1986).

[22] For example, Frenkel (1978, 1980, and 1981b) suggested that the real exchange rate is nonstationary and was not aware that his results were spurious.

[23] Notice that if the researcher's null hypothesis is stationarity, then classical statistics will be appropriate. Thus, when testing whether a series exhibits a stationary process or not, one might get a contradiction if he once assumes the null is stationarity and once the null is non-stationarity.

horizon, but for any given finite sample there exists a representative from either class of stationarity that could account for all the features of the data. Thus, several recent papers have argued that any attempt to identify the true generating process is inherently impossible and misleading (see Blough, 1992, Cochrane, 1991, Christiano and Eichenbaum, 1990, and Stock, 1990). Thus the researcher must be cautious with his inferences and conclusions and should employ several alternative procedures to confirm his findings. Furthermore, the separate estimations of eq. (7), (8), and (9) do not exploit the cross information and do not distinguish between short and long run dynamics, thereby having low power.

The above discussion strongly supports the use of CVECM analysis. The formal CVECM analysis consists of three steps: (1) testing for stationarity (2) testing for cointegration and (3) estimating the ECM (if nonstationarity and cointegration are found). At each step, several advanced statistics must be used that account for nonstationarity and cointegration. Further, the inclusion of dummy and stationary variables must be allowed for in the ECM.[24]

There are at least three theoretical and three empirical advantages of using CVECM analysis. First, on the theoretical level, CVECM is applicable to variables which are nonstationary but still connected in the steady state. Thus, there exists a stationary linear combination of the variables which evolves around its mean and those deviations from this mean taper off to zero in the long run. This is the exact pattern predicted by PPP theory and therefore, CVECM is a natural framework for this kind of time series analysis. The CVECM tests stationarity vs. nonstationarity for which is a test of the PPP model vs. Balassa and random walk models. In the case of nonstationarity, it goes on to test for the existence of deterministic trend and stochastic drift as proposed by the latter models (allowing for the simultaneous existence of both- "pooling") while testing the PPP's validity after subtracting the trend, drift and fundamentals. Moreover, it is able to

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[24] I use the word "complete," since I use the complete package of the VECM and fully exploit it, especially by using trend and/or drift nonstationary variables, stationary variables, dummy variables and by employing several test procedures in each step of the analysis.

distinguishing between short and long run real effects and dynamics. Finally, it has the ability to test PPP in the presence of transportation costs and tariffs. It is also worthwhile noting that it bypasses the index number problem mentioned at the end of section 2 and removes some of the trend effects common in time series.

On the empirical level, the first attraction of the analysis is the simplicity of the estimation which uses OLS. Second, all the variables can be treated as jointly endogenous, which is necessary for the application of OLS to PPP and which avoids the inconsistency discussed by Davutyan and Pippenger (1985), Edison (1985), Frenkel (1978), and Krugman (1978).[25] Third, there is no need to assume serially uncorrelated innovations.

In spite of the above attractions, one should keep in mind the above caution remark regarding the problems embodied in any endeavour to identify the true generating process, although this approach uses advance statistics and methodology. Farther, different inferences may emerge from the different statistics, and researcher must reflect and approach such results fair enough.

The following sub-sections present the steps of the CVECM, the statistics which shall be employed in each step and a brief discussion of their pros and cons.

4.1. Testing for stationarity[26]

It should be emphasised that while researchers have tested for stochastic drift nonstationarity, they have never tested for deterministic trend nonstationarity as an additional or opposing hypothesis. This paper will provide such a test.

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[25] The aforementioned authors pointed out that the OLS results may be inconsistent due to simultaneous equation bias, i.e. both exchange rate and prices are endogenous.

[26] A number of other approaches for unit root testing have been proposed (see Stock (1993) for an excellent survey). The approach adopted in this research is widely used and has a good track record.

I present four asymptotically equivalent procedures for detecting nonstationarity, all of which allow for fitted drift and time trend and take into account serial correlation and potential heteroskedasticity in the disturbances. The first procedure engages a visual examination of the graph of the series itself, the sample correlogram and the sample partial correlogram. For stationarity, the correlogram should be smaller than one (in absolute value) and die away rapidly as the lag increases. However, this test cannot determine whether the series has a trend and/or a unit root. It is odd that, to my knowledge, only Edison (1985) Froot et al. (1995), and Phylaktis (1992) have executed visual tests of the plotted series and furthermore, no paper has tested the correlogram and partial correlogram.

The second procedure is the augmented Dickey-Fuller ADF-test.[27] Said and Dickey (1984) have shown that an unknown ARIMA(p,d,q) process can be well approximated by an ARIMA(p,d,0), where $p \leq T^{1/3}$. Therefore, they argue that the ADF test is approximate for economic time series containing autoregressive and moving average components of unknown order when applied for approximated ARIMA(p,d,0). Thus, I begin by applying the Box and Jenkins (1976) method to identify a tentative ARIMA(p,d,0) model with parsimonious representation for each of the series as in the following general auxiliary equation:

$$\textbf{(10)}\ \Delta^d y_t = \beta + \sum_{i=1}^{n} \kappa_i D_{it} + \delta t + \rho \Delta^{d-1} y_{t-1} + \sum_{j=1}^{p} \zeta_j \Delta^d y_{t-j} + \varepsilon_t\,;$$

where y represents the level for each of the series, D_i is a dummy variable represents seasonality and structural change, t is the time index and ε is a white noise.

Though it is very important for the ADF-test not to have correlation present in the residuals, consequent inclusion of insignificant augmentations would reduce the power of the test to reject the null of a unit root because a reduction in the degree of freedom.

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[27] The ADF-test is due to Dickey and Fuller (1979, 1981) and modified by Bhargava (1986).

Therefore, following the suggestion by Said and Dickey (1984), I drop lagged terms that are jointly insignificant by conventional F-tests at the 10 percent level, beginning with the longest lag of 6. If omission of insignificant lags produces serial correlation in the residuals, as indicated by the Ljung Box Q_{LB}-statistic *or* the Breusch-Godfrey Lagrange Multiplier L_{BG}-statistic at the 10 percent level, then lags are reintroduced to eliminate the problem.[28]

Since more than one unit root is hypothesized, I conduct a sequence of tests as recommended by Dickey and Pantula (1987) beginning with two (largest) hypothesized unit roots.[29] Most papers published so far have mistakenly tested for only the first order of integration.

It is worth noting that differencing a seasonal unit root will not yield a stationary process. Further, when the DGP incorporates a structural change, then the ADF-test is biased toward the non-rejection of a unit root. Perron (1989) has extended the ADF-test to allow for *one-time known* change in the constant and the trend by including dummy variables, as is shown in eq. (10). Further, if one suspect for a deterministic seasonality exists, then he should deseasonalize the series before testing stationarity by running the following regression:

$$\textbf{(11)}\ \tilde{y}_t = \kappa_0 + \sum_{i=1}^{n} \kappa_i D_{it} + y_t \ ;$$

where D_i is the seasonal dummy, $\tilde{y}$ is original seasonal series and y is the residuals which represents the depersonalized series. Thereafter, use y for eq. (10) without the dummy variables and apply the ADF-test, which will still be possible as Dickey *et al.*,

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[28] In case of doubt, the Von Neumann Ratio (VNR) was also applied. Note that the above statistics are superior to the Durbin-Watson d_{DW}- and h_{DW}-statistics and Box-Pierce Q_{BP}-statistic. Further, the L_{BG}-statistic seems to be more conservative, in the sense that if it rejected the null hypothesis that there is serial correlation when the Q_{LB}-statistic rejected it, but not visa versa.

[29] Notice that the opposite direction, which starts from the smallest hypothesized unit root, tends to select too few unit roots.

1986) have shown. This extension can be useful in examining the seasonal, seasonal unit root, and structural and regime change.[30]

Finally, I follow the suggestion of Doldado *et al.* (1990) and apply ADF-test for a sequence of null hypotheses as described in Table . First I check for the significance of the trend and the null of unit root, and then I test the significance of the drift and the null of a unit root, while choosing the appropriate tabulated critical values.

Although the ADF-test is widely used, one should keep in mind that its power is limited. As has been mentioned, the test has a low power to reject the null I(1); hence, it only allows us to reject (or fail to reject) the null hypothesis of I(1). Therefore, a rejection of the null I(1) is a strong evidence in favor of I(o) and one might not proceed farther to support his findings, but a failure to reject (especially at a high significance level) is only weak evidence in favor of the I(1) null and further investigation should be held. Since there may be a difference in the small sample performance of DF- and ADF-statistics (see Engle and Granger, 1987), one must also be cautious with regard to non-rejection of the I(1) null. In addition, while the ADF-test is an improvement over the traditional tests, it has low power in detecting near unity. Moreover, Choi (1990), quoted by Green (1993) (p. 563), has suggested that the ADF-test is distorted by the presence of moving average components in the innovations. It is worth noting that Dickey and Fuller (1979, 1981) indicate that the Z_{DF} test is more powerful than the t_{DF} test. Further, if the true DGP contains a drift or trend, then if the sample size big enough, one should use the regular standard normal distribution (see Campbell and Perron, 1991, and Doldado *et al.*, 1990).

The third procedure is the Philips and Perron PP-test, which is robust to a wide variety of heterogeneously distributed and serially correlated innovations.[31] The findings from the above stage, which indicate heterogeneity and dependence in the innovation

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[30] For farther consulting see the *Journal of Business and Economic Statistics*, July 1992.

[31] The PP-test is due to Perron (1988), Philips (1987), and Philips and Perron (1988).

sequence, motivate the use of this test. The test involves computing the following OLS regression:

$$\textbf{(12)}\ \Delta y_t = \beta + \rho y_{t-1} + \delta t + u_t .$$

where y represents the level for each of the series, t is the time and u has mean zero but can be heterogeneously distributed and serially correlated.

The PP-test provides three statistics, t_{PP}, Z_{PP} and $\phi_{(3)DF}$, and I use them all to test the same null hypotheses which appears above. The reported critical values are the same as those mentioned above for the ADF-test. Further, this test allows inclusion of dummy variables for seasonality and structural change, as has been explained above for the ADF-test, thereby increasing its precise estimation of the integration order.

The PP-test has a greater power to reject a false null hypothesis of a unit root. It also allows, as mentioned above, for a weaker assumption concerning the error term. However, the test must be applied with caution if MA terms in the ARIMA model are negative, since in this case the test tends to reject the null of a unit root, whether or not the true DGP contains a negative unit root.[32] The above criticism concerning the ADF-test of unit root applies also for the PP-test.

The fourth procedure is the Johansen J-test.[33] This procedure tests the null hypothesis of cointegration for a VAR system, by estimating Full Information Maximum Likelihood (FIML). Applying this procedure for testing the hypothesis of no-cointegration within a univariate "VAR" is equivalent for testing the hypothesis of no unit root. Further, if the J-test shows that the cointegration rank equals the number of the variables in the VAR, then it means that non of the series has a unit root, and this must not contradict the findings of the ADF-test.

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[32] For further criticism, see Banerjee *et al.* (1993), pp. 109-119, Enders (1995), p.242, and Hamilton (1994), pp. 515-516.

[33] The test is due to Johansen (1988, 1991).

Note that R^2 cannot be used as a criterion; however, one can use instead the Akaike Information Criterion (AIC) and the Baye Information Criterion (SBC). These criteria are used for support but not as primary tools. They place a penalty on extra coefficients. The model with the minimum AIC- or SBC-statistics will be chosen. Hannan (1980) has generalized these criteria and determined conditions under which they are consistent in choosing the "correct" model. However, the evidence presented in Sneek (1984) suggests that there is a tendency for the AIC to choose an over-parameterized model. Nevertheless, in some specifications (e.g., white noise with same number of parameters), the R^2 and AIC select the same model.

4.2. Testing for cointegration

A vector of variables (x_t) is said to be cointegrated if each of the variables has a stationary inevitable ARIMA representation after being differenced d times (x_t~I(d)), and are tied together in one long run relationship, i.e. there exists a long run linear combination I(d-b) (d≥b>0) of the variables (x_t~CI(d,b)).[34] We are interested in the case d=b (and not necessarily equal to one, as some researchers have claimed), i.e. deviations from the long run equilibrium combinations (also called long run residuals or the long run error term) must tend to damp to zero in the long run. In econometric terms, it simply means that a unit root of the long run residuals is ruled out.

The cointegration relation is expressed and estimated by the following VAR model:

$$(13)\ \alpha x_t = c_0 + \delta t + \phi s d_t + \upsilon_t;$$

where $\alpha = \begin{pmatrix} 1, \alpha_{12}, \alpha_{13} \\ \alpha_{21}, 1, \alpha_{23} \\ \alpha_{31}, \alpha_{32}, 1 \end{pmatrix}$ is the cointegration matrix (CIM), $x_t = \begin{pmatrix} s_t \\ -p_t \\ p_t^* \end{pmatrix}$, $c_0 = \begin{pmatrix} c_{01} \\ c_{02} \\ c_{03} \end{pmatrix}$ is the vector of the real exchange rate minus transaction costs, $\phi = \begin{pmatrix} \phi_{11}, \phi_{12}, \cdots, \phi_{1f} \\ \vdots \\ \phi_{f1}, \phi_{f2}, \cdots, \phi_{ff} \end{pmatrix}$,

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[34] Concepts and tests of cointegration are presented in Engle and Granger (1987).

$sd_t = \begin{pmatrix} sd_{1t} \\ \vdots \\ sd_{ft} \end{pmatrix}$ is any other possible structural stationary and dummy variables and

$\upsilon_t = \begin{pmatrix} \upsilon_{1t} \\ \upsilon_{2t} \\ \upsilon_{3t} \end{pmatrix}$ is the vector of the long run residuals. The right-hand-side of Eq.

(13) is not affected whether any of the variables exhibits a non-zero drift, however, it includes a trend to cover the possibility that one variable or more exhibit a non-zero trend.

Eq. (13) implicitly assumes and imposes that the coefficient of the dependent variable in the cointegrating vector (CIV) is non-zero. Thus if cointegration exists and CIV does not have a non-zero coefficient for the normalized variable, then arbitrary selection of the normalized series will result in a fundamentally misspecified model, and will materially affect the estimate of the CIV as well as the evidence for cointegration among the series. Therefore, it is important to estimate the three regressions in eq. (13) such that each series will be the normalized dependent variables; thereafter cointegration in all the regressions is tested for. Another important reason for estimating eq. (13) is that if the normalized CIV of each of the regressions is not the same, then this is evidence for the existence of more than one CIV. In other words, the rank of the CIM gives the number of the CIVs. Many researchers had not been cautious on these points, which probably have affected their findings and conclusions.

Note that other than for the case of a pair of series, the series do not need to be integrated of the same order as many researchers have argued. If they are not integrated of the same order, then there exists a more sophisticated inter-relationship among the series and further explanation of this inter-relation is required.[35]

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[35] Consider, for instance, the case where e and p are I(1) and p* and v are I(0). Since v is I(0), integration exists. One interpretation of this case is that the domestic country is a small

The possibility of including a trend, stationary, and dummy variables is crucial for my complete nesting methodology. This tool enables testing a structural models and a change in the regime, such that the significance of the demand and supply fundamentals and even nominal rigidities and pricing to market. Moreover, it enables testing whether subtraction of these factors will reveal PPP or whether there exist other factors which should be considered exploring the structural model of the deviations from PPP.

Under cointegration, the CIM can essentially be estimated by OLS, irregardless of the endogeneity of the regressors with respect to the long run error term (Engle and Granger, 1987)). This is because any correlation between them will be of a lower order in T [the number of observations used in the estimation] than the variance of the regressor itself. Further, the OLS estimator of the CIM is consistent (with bias O(1/T)), highly efficient and converges to the true population value at a rate even greater than in the standard regression (with variance $O(1/T^2)$ as opposed to the standard case of O(1/T)). Therefore, the standard asymptotic theory does not apply and therefore one cannot use the reported *t*-statistics for valid inference (see Banerjee et al., 1986, and Stock, 1987)).[36]

West (1988) has modified the standard *t*-statistic so that it yields an asymptotically normal statistic and can therefore be employed in making statistical inferences about the estimated coefficients of I(1) regressors.

The first procedure for testing cointegration involves testing the stationarity of the long run residuals as defined in eq. (13) by applying the visual test, ADF-test and the Philips and Ouliaris PO-test (or Philips and Hansen PH-test if at least one of the variables

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market and the PPP is maintained by adjustments in that market. Moreover, under a fixed exchange rate regime, one might expect $(p,p^*)\sim I(1)$ and $(e,v)\sim I(0)$ due to the fact that revision towards PPP occurs through nominal prices rather than the nominal exchange rate.

[36] Many researchers, who have utilized the semi-ECM of the above framework, were not aware of this problem (see, for instance, Adler and Lehmann, 1983, Apte *et al.*, 1994, and Edison, 1985, 1987).

exhibits a non-zero drift or trend) without trend and constant.[37] However, when neither of the variables exhibits a non-zero drift or trend then the critical value of DF for testing stationarity by using ADF-test and PO-test does not hold for the *estimated* residuals since they do not take in account that the series of the residuals is not the *true* series. Critical values which take this fact into account are reported by Engle and Granger (1987) and Phillips and Ouliaris (1990) and expanded by Engle and Yoo (1987) and Haug (1992) for small samples.[38] If at least one of the variables exhibits a non-zero drift or trend then when using PH-test the critical value of DF will still hold (basically because the drift or the trend dominate the asymptotic distribution of the residuals). A complementary test which accounts for this fact is the Sargant-Bhargava SB-test.[39] This test involves comparing the Durbin-Watson d_{DW}-statistic with critical values tabulated by Engle and Granger (1987) and extended by Engle and Yoo (1987) for small samples. Engle and Granger (1987) have demonstrated its high power.

The second procedure involves testing the significance of the long run error term in the ECM (see below) by applying the Wald test and testing whether it is bounded between zero and -2. Kremers et al. (1992) has shown this test of cointegration to be more powerful than the ADF test on the long run residuals. The distribution of this statistic lies somewhere between a $N(0,1)$ and DF distribution. The most conservative values will be those from the DF distribution.

The third procedure is the one step FIML of the Johansen J-test. This approach avoids the normalization problem. Further, this approach is useful in testing for the possibility of more than one cointegration vector (cointegrating rank greater than one). Also, it simultaneously estimates the short and long run dynamics (one step ECM estimation),

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[37] The PO-test is due to Philips and Ouliaris (1990), and PH-test is due to Hansen (1992).

[38] The difference between these critical values and the DF critical values is that the former examine the second largest eigenvalues in the bivariate time series, while the latter examine the first largest eigenvalues in the univariate time series. Therefore, the former have to be raised in order to correct the test bias.

[39] The test is due to Sargan and Bhargava (1983).

thereby helping to overcome the small sample bias.[40] The estimators are thus more efficient and the J-test for cointegration has more power than the above tests.[41]

A fourth possible procedure is the Park P-test which tests the null of cointegration. This test is constructed by adding a time polynomial to the right-hand-side of eq. (12) and testing its significance.

4.3. Estimating the ECM

The ECM estimation is carried out by estimating the long run relation as expressed in eq. (13) (which already has been done in the second step above) and the short run dynamic relation as expressed by the following parimonious model (smallest h) estimated by OLS:

$$\textbf{(14)}\ \psi\Delta x_t = c_{1t} + \varphi \upsilon_{t-1} + \psi^1 \Delta x_{t-1} + \psi^2 \Delta x_{t-2} + \ldots + \psi^h \Delta x_{t-h} + u_t\,;$$

where $\psi = \begin{pmatrix} 1, \psi_{12}, \psi_{13} \\ \psi_{21}, 1, \psi_{23} \\ \psi_{31}, \psi_{32}, 1 \end{pmatrix}, \psi^i = \begin{pmatrix} \psi^i_{11}, \psi^i_{12}, \psi^i_{13} \\ \psi^i_{21}, \psi^i_{22}, \psi^i_{23} \\ \psi^i_{31}, \psi^i_{32}, \psi^i_{33} \end{pmatrix}, \varphi_t = \begin{pmatrix} \phi_{1t} \\ \varphi_{2t} \\ \varphi_{3t} \end{pmatrix}, u_t = \begin{pmatrix} u_{1t} \\ u_{2t} \\ u_{3t} \end{pmatrix}$ is the short run residuals and x_t and v_t are as defined in eq.(13). The parsimoniously of the model dtermined by droping elements that are jointly insignificant by conventional *F*-statistic and L_{BG}-statistic at the 10 percent level, and by a support of the SBC. The distribution of this statistic lies somewhere between an *N*(0,1) and *DF* distribution. The most conservative values will be those from the *DF* distribution.

I estimate the ECM using the one step Johansen procedure and the two step Engel-Granger procedure. After I have estimated the final representative econometric model for the exchange rate process, I then turn to testing the 'proportionality' null hypothesis:

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[40] For another related approach, see Phillips and Hansen (1990) and Park (1992).

[41] Horvath and Watson (1993) have extended the J-test to allow for constraints that represent long run equilibrium conditions (this effectively transforms the J-test into a two step ECM). For applications see Edison et al. (1994).

$$H_o^f : \alpha = I\_\&\_c_1 = 0;$$

which implies the existence of the PPP. I further execute an 'asymmetry' diagnostic, by estimating the following (bivariate system) regression:

$$\textbf{(15)}\ s_t = \bar{c} + \bar{\alpha}(p_t - p_t^*) + \vartheta_t,$$

and testing the null hypothesis:

$$H_o^g : \bar{\alpha} = 1\_\&\_\vartheta_t \sim I(0).$$

Further possibility to test the PPP, is to impose the proportionality and test its significance and the stationarity of the residuals. In other words, to estimate the following equation:

$$\textbf{(16)}\ e_t = \overset{=}{c} + \vartheta_t,$$

and testing the following null hypothesis:

$$H_o^h : \vartheta_t \sim I(0).$$

5. The empirical findings

During the 1920s a floating exchange rate regime prevailed between the US and the UK. This era has been studied by Clement and Frenkel (1980, 1981), Davutyan and Pippenger (1985), Edison (1989), Frenkel (1976, 1978, 1980, 1986), MacDonald (1985), Phylaktis (1992), and Taylor and McMahon (1988). However, as stated at the beginning of this article, no consensus has emerged. This section will re-examine this period and try to shed some light. I use same database that has been used by the above research in order to maintain comparability (details regarding the data are given in Appendix 1). Though this is an implication of the proposed methodology for testing the robustness of different exchange rate models, I only test the PPP.

5.1. Testing for stationarity[42]

Graph 1 reveals some clear trends in the series of the $/£ exchange rate and the UK-WPI and to a lesser extent in the US-WPI series. Nevertheless, a quick glance at the sample correlograms and partial-correlograms (Graph 2) reveals quite unambiguous evidence for the nonstationarity of the three series: (i) the three sample correlograms gradually decay toward zero while the sample partial-correlograms fall immediately after the first lag indicating an AR(1) process; (ii) in an AR(1) process the first autocorrelation equals the coefficient of the first lag, which is near unity in our case, suggesting the strong possibility of a random walk or time trend.[43]

ADF-test and PP-test for the estimated regression are reported in Tables 1 and 2 and the statistics are reported in Table 3. The results according to both tests reveal convincing evidence for the nonstationarity of the three series. The null hypothesis of a unit root ($H_o^a : -2 < \rho < 0$) cannot be rejected in all cases with even 10 percent significance level (1 percent significance level when using t_{PP} for the case of UK-WPI). The trend is insignificant at most cases with even 10 percent level, and some times 1 percent level, implying a rejection of the null $H_o^b : \delta = 0$. The joint null hypothesis of a unit root without a trend ($H_o^c : -2 < \rho < 0\_\&\_\delta = 0$) cannot be rejected even with 10 percent significance level. Since the findings are highly convincing, there is no need to use further tools. Notice also that the drift is significant in all cases.[44]

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[42] I began by checking whether the series exhibit I(2). The results were remarkably unambiguous indicating a stationary first difference. Due to limitation of space, only the results of the test for stationarity of the levels are reported.

[43] For an explanation of the detection and identification of the stationarity of the series, see Enders (1995), Ch. 2, and Pindyck and Rubinfeld (1991), Part 3.

[44] Following the suggestion of Doldado *et al.* (1990), because the trend is insignificant, the test of stationarity was redone by regressing eq. (10) without a trend, and the unit root was still not rejected. Hence, I tested the significance of the drift separately and jointly with the unit root. I found that the drift is significance. However, I used the conservative DF critical value and not the standard normal distribution as they suggest in this case, since I was concerned of a small

Though "findings are highly convincing", one might discard them by the claim that these results are misleading due to a sharp adjustment during the first year after abandon the Gold Standard and during the first year before restoring the Gold Standard. However, when applying ADF-test for the sample period February 1922 through May 1924, I still cannot reject the null unit root.[45] Further, I do not believe and there is no indication (e.g. from the correlograms or the series) for the existence of seasonality, and the sample size does not permit testing this. Therefore, I do not include dummy variables.

It is worthwhile noting that the attempt of Taylor and McMahon (1988) to test stationarity when subtracting the last year seems to be imprecise, since it substantially shortens the period such that it does not cover enough cycles. The appropriate technique for testing a structural change is not by shortening and change the sample size, but by applying Perron's test. Nevertheless, when applying the above technique I still cannot reject the null unit root.

Since I conclude that the series are I(1), I proceed to test for cointegration in the next section.

5.2. Testing for Cointegration

In order to test CI, I estimate eq. (13). The estimated cointegration matrix is reported in Table 4. Then I test the stationarity of the residuals vector υ_t. Graph 3 reveals evidence in favor of the nonstationarity of the estimated residuals vector. The correlograms damps geometrically and the partial correlograms damps immediately after the first lag, while the first autocorrelation is near unity. A support for this conclusion comes from the statistics reported in Table 6. The unit root hypothesis of the residual series is not rejected, implying the inability to reject the hypothesis of no cointegration; the J-test

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sample bias and also because the visual test suggests that the series most likely exhibit either a trend or a drift. Due to limitation of space, these results were not reported.

[45] The t_{DF} is -1.7, -2.4 and -1.9 for the \$/£ exchange rate, US-WPI and UK-WPI which is greater than the critical value of -2.97 (-3.68) at 5 (10) per cent significance.

results support this conclusion. However, the long run error terms in the ECM are bounded between zero and -1, and significant when the exchange rate and US-WPI are normalized, but the CIV in this case is far from unity.

5.3. *Estimating the ECM*

Though the previous section suggests of no cointegration, I proceed to estimate the ECM for the purpose of suggesting another procedure for testing cointegration as already has been used above. This estimation does support the existence of cointegration and that c_1=o, but the CIV does not seem to be unity, which means the $H_o^f: \alpha = I\_\&\_c_1 = 0$ is rejected.

I proceed to check the asymmetry hypothesis by estimating eq. (15):

$$s_t = \underset{(0.026)}{1.841} + \underset{(0.049)}{0.687} (p_t - p_t^*)$$

which indicates that asymmetry hypothesis ($H_o^g: \overline{\alpha} = 1\_\&\_\vartheta_t \sim I(0)$) is also rejected.

Further, I impose the unity CIV and test whether it reveals a stationary residuals. To do so, I estimate eq. (16):

$$e_t = \underset{(0.006)}{2.00}$$

and test whether the residuals has a unit root. The t_{DF}=-2.574, which is greater than -3.27 (-2.99) as the 5 (10) percent significant level. Graph 4 shows that the revision to the PPP takes more that half of the sample period, and it took place only once, which means there is no tendency for revision and casts real doubts about the PPP. Graph 5 which plots the correlogram and the partial correlogram also indicate that residuals exhibit at least a near unity. Therefore, $H_o^h: \vartheta_t \sim I(1)$ cannot be rejected implying inability to reject the null of no PPP.

However, to be more precise, I emphasize that the findings cannot reject the null of invalidity of the PPP, which does not mean a rejection of the PPP. It is rather weak evidence against the PPP. In this stand, we must also recall that the consensus emerged from the papers that support strong evidence in favor of the PPP that there is in fact a moderate tendency for real exchange rates to converge towards long run equilibrium. The half-life of PPP deviation appears to be four years - 15 percent annual revision (see Froot and Rogoff, 1994, who provides an excellent critical survey and review of this literature, see also Frankel and Rose, 1995). Therefore, it is hard to be able to support the PPP theory or even fairly reject it when employing only about five and half year of sample period. The some how little evidence that has been found in this paper in favor of the PPP and the doubt to reject it reflect this fact. I guess the most supportive evidence for the attitude can be obtained from inspection of Graph 4.

6. Summary and conclusions

The paper proposes a methodology that makes use of advance techniques for testing the robustness of the exchange rate models. These techniques proved to be powerful and appropriate for testing the robustness of the various exchange rate models and their pooling version. Besides its simplicity, the methodology accounts for the various features that might describe the exchange rate and the price indices.

When applying this methodology for the 1920s experience between the US and UK, the null of the invalidity of the PPP cannot be rejected. This conclusion cannot be said to concur with evidence generated by MacDonald (1985) and Edison (1985), neither to be at variance with the results of Clement and Frenkel (1980, 1981), Davutyan and Pippenger (1985), Frenkel (1976, 1978, 1980, 1986), and Taylor and McMahon (1988). Other than the imprecise of the econometric application, the main reason for the odds is my distinction between rejection the null of PPP and inability to reject the null of invalidity of PPP. Further, I recognize the limit of the econometric instruments to test the robustness of the PPP by employing only four years sample period.

Balassa (1964) claims that there was a maladjustment for the $/£ exchange rate in 1925. Graph 4 implies that $/£ exchange rate was under adjusted about 6 percent relatively to its purchasing power at the time when UK abandoned the Gold Standard. However, this is above the average value for the floating period, and purchasing power was calculated relatively to the WPI. We might get a different answer if we employ a different index. Another worth mentioning observation is that the price index of the US was not responding for any deviation from the PPP, whilst the UK-WPI did respond, which might be interpreted that the UK economy was a small one relatively to the US economy. Have we believed that there is a cointegration, then a more formal test would be to use the impulse respond function.

While this paper has highlighted some of the inadequacies of the econometric analysis that had been applied for the PPP verification and the invalidity of the PPP for the 1920s floating, its main contribution is a positive one: it suggests a precise and simple methodology for a joint test of the robustness of the exchange rate models and their pooling model. This methodology can be applied for testing other stylized facts in macro economics. Another point is to argue that econometrist must be aware of the limit of the econometric techniques. The developing of these techniques will continue in the future and might bring another attitude that changes our judgment about the robustness of the exchange rate models.

Table 1: The null hypothesis that are tested for determining the class of stationarity of the series

The null hypothesis		Statistics employed	Tabulated critical values used
Unit root-	$H_o^a{:}{-2} < \rho < 0$	*t*- and *Z*-statistics	Dickey (1976), Dickey and Fuller (1989), Fuller (1976), MacKinnon (1991), Phillips (1987), and Perron (1989)
No deterministic trend-	$H_o^b{:}\delta = 0$	*F*, L_M and *t*-statistics	Regular *F*, X^2 and *t* tables
Unit root and no trend-	$H_o^c{:}{-2} < \rho < 0\_\&\_\delta = 0$	$\phi_{(3)}$-statistic	Dickey and Fuller (1981)
No drift-	$H_o^d{:}\beta = 0$		Dickey (1976), Dickey and Fuller (1989), Fuller (1976), MacKinnon (1991), Phillips (1987), and Perron (1989)
Unit root and no drift-	$H_0^e{:}\beta = 0\_\&\_\rho = 0$		Dickey and Fuller (1981)

Note: It should be emphasize that a two-tailed test should be held for testing stationarity rather than one-tailed as always done in literature. Fuller (1976) has derived the appropriate limit distributions, Dickey (1976) has computed empirical approximations; both were published in Dickey and Fuller (1979). MacKinnon (1991) has derived a formula which calculate the DF critical values for any sample size and any of the choices of right-hand-side variables. Phillips (1987) has replicated these calculations for the PP-test (see below). Perron (1989) has calculated the critical values when a dummy variable is included. Notice that for multi-exchange rates, I would need to find critical values by calculating the exact small sample distribution via Monte Carlo estimates or some other procedure.

The table follows the suggestion of Doldado *et al.* (1990) by conducting a sequence of diagnostic test of the null unit root. First the following equation is estimated: $(10)\,\Delta^d y_t = \beta + \sum_{i=1}^{n} \kappa_i D_{it} + \delta t + \rho \Delta^{d-1} y_t + \sum_{j=1}^{p} \zeta_j \Delta^d y_t + \varepsilon_t$; where y represents the level for each of the series, D_i is a dummy variable represents seasonality and structural change, t is the time and ε is a white noise. Then I test the null of unit root by using t_{DF}, t_{PP}, Z_{DF} and Z_{PP} statistics and employing the DF critical values. If the null was rejected then I conclude the series is stationary and no further investigation is done. Otherwise, I test the significance of the regressors using *F, L_m and t* statistics and employing the regular tables. In particularly, I test the null no-trend and the joint null of unit root and no-trend by using the t and $\phi_{(3)}$ and the DF critical values. If I cannot reject the null no-trend then I proceed to test the null unit root by applying regular tables. Otherwise, I regressing eq. (10) without the trend and test the null unit root by applying t_{DF}, t_{PP}, Z_{DF} and Z_{PP} statistics and employing the DF critical values. If it is not rejected then I proceed to test the null no-drift by using t_{DF}, t_{PP}, Z_{DF}, Z_{PP} and $\phi_{(1)}$ statistics and DF critical values. If I cannot reject the null no-drift then I retest the null unit root by applying regular tables. Otherwise, I regressing eq. (10) without the trend neither the drift, and test the null unit root by applying t_{DF}, t_{PP}, Z_{DF} and Z_{PP} statistics and employing the DF critical values.

Table 2: The results of the estimated regressions for testing the stationarity of the series by applying ADF- and PP-tests

	β	δ	ρ	ζ_1	ζ_2	#OBS.
ADF-test						
s	**0.182**	**0.00048**	**-0.131**	**0.193**	**0.360**	49
	(0.076363)	(0.000273)	(0.055426)	(0.137849)	(0.140281)	
p	**0.787**	**0.00029**	**-0.173**	**0.434**	-	**50**
	(0.254440)	(0.000141)	(0.055877)	(0.119812)	-	
p*	**0.327**	**0.00004**	**-0.065**	**0.535**	**-0.038**	**49**
	(0.132580)	(0.000138)	(0.025840)	(0.140949)	(0.136669)	
PP-test						
s	**0.115002**	**0.000258**	**-0.079874**	-	-	**51**
	(0.076811)	(0.000277)	(0.055728)	-	-	
p	**0.770451**	**0.000401**	**-0.170448**	-	-	**51**
	(0.259741)	(0.000141)	(0.056949)	-	-	
p*	**0.550221**	**0.000230**	**-0.109908**	-	-	**51**
	(0.124666)	(0.000140)	(0.023938)	-	-	

Notes: s is the \$/£ nominal exchange rate, p and p* are the US-WPI and UK-WPI, respectively. ADF-test and PP-test denote augmented Dickey-Fuller and Phillips-Perron tests, respectively. The estimated regressions are:

For the ADF-tests: $(10)\, \Delta y_t = \beta + \delta t + \rho y_{t-1} + \zeta_1 \Delta y_{t-1} + \zeta_2 \Delta y_{t-2} + \ldots + \zeta_{p-1} \Delta y_{t-p+1} + \varepsilon_t ;$

For the PP-tests: $(12)\, \Delta y_t = \beta + \rho y_{t-1} + \delta t + u_t ;$

where y represents any one of the above series, ε is a white noise, and u can exhibits a serial correlation. Standard errors appear in parentheses. The sample period is February 1921 through May 1925, during which a floating \$/£ exchange rate regime prevailed.

Table 3: The results of the statistics for testing the stationarity of the series by applying ADF- and PP-tests

		H_0: ρ=0				H_0: δ=0			H_0: ρ=0 & δ=0
		PP-test		ADF-test		ADF-test			ADF-test
		t_{PP}	Z_{PP}	t_{DF}	Z_{DF}	F	L_M	t	ϕ_3
s		-1.75	-6.08	-2.37	-14.37	3.13	3.36	1.77 ①	2.81
p		-3.10	-10.40	-3.10	-15.33	4.26 ①	4.43	2.06 ①	5.04
p*		-3.81 ①	-6.12	-2.50	6.30	0.10	0.11	0.31	3.39
Critical values:	1%	-4.15	-25.7	-4.15	-25.7	5.08	9.21	2.40	9.31
	5%	-3.50	-19.8	-3.50	-19.8	3.19	5.99	1.68	6.73
	10%	-3.33	-16.8	-3.33	-16.8	-	4.61	1.30	5.61
Table		B.6(4)	B.5(4)	B.6(4)	B.5(4)	B.4	B.2	B.3	B.7(4)

Notes: s is the $/£ nominal exchange rate, p and p* are the WPI of the US and UK, respectively. ADF-test and PP-test denote augmented Dickey-Fuller and Phillips-Perron tests, respectively. The sample period is February 1921 through May 1925. The estimated statistics outlined by ①, ⑤ and ⑩ indicate significance at 1, 5 and 10 percent levels, respectively. The critical values for t_{DF} and t_{PP} were calculated directly by MacKinnon formula. The critical values for Z_{DF} and Z_{PP} are tabulated in Dickey and Fuller (1979) and Phillips (1987). The critical values for F, L_M and t are the regular tables. The critical values for ϕ_3 are taken from Dickey and Fuller (1981). For convenience, the last row indicates the appropriate table in Hamilton (1994). The table presents unambiguous evidence in favor of the nonstationarity of the series.

Table 4: The long run dynamic- the cointegration regression

	c	s	p	p*
Two step Engle-Granger procedure				
s	0.057	-	1.001	-0.620
p	1.967	0.582	-	0.344
p*	1.613	-1.206	1.152	-

Notes: *s* is the $/£ nominal exchange rate, p and p* are the WPI of the US and UK, respectively. The sample period is February 1921 through May 1925. Standard errors are not reported since they may be misleading in this context. A trend is not included since neither of the variables exhibits a trend. The one step Johansen procedure was not used since the Johansen test cannot reject the hypothesis of no cointegration. The CIV are not proportional to unity, except for the case when the UK-WPI is normalized.

Table 5: The results of the estimated regressions for testing the stationarity of the residuals from the cointegration regression by applying ADF- and PH-tests

	ρ	ζ_1	ζ_2	#OBS.
ADF-test				
υ_s	**-0.244**	**0.086**	**0.317**	**49**
	(0.091)	(0.142)	(0.140)	
υ_P	**-0.246**	**0.125**	**0.30**	**49**
	(0.085)	(0.132)	(0.135)	
υ_{P*}	**-0.212**	**0.053**	**0.282**	**49**
	(0.078)	(0.139)	(0.135)	
PH-test				
υ_s	**-0.176**	-	-	**51**
	(0.084)	-	-	
υ_P	**-0.196**	-	-	**51**
	(0.080)	-	-	
υ_{P*}	**-0.164**	-	-	**51**
	(0.072)	-	-	

Notes: υ_s is the residuals when the $/£ nominal exchange rate is normalized, υ_P and υ_{P*} are the residuals when the US-WPI and UK-WPI are normalized, respectively. ADF-test and PH-test denote augmented Dickey-Fuller and Phillips-Hansen tests, respectively. The estimated regressions are:

For the ADF-tests: (10) $\Delta y_t = \beta + \delta t + \rho y_{t-1} + \zeta_1 \Delta y_{t-1} + \zeta_2 \Delta y_{t-2} + \ldots + \zeta_{p-1} \Delta y_{t-p+1} + \varepsilon_t$;

For the PH-tests: (12) $\Delta y_t = \beta + \rho y_{t-1} + \delta t + u_t$;

where y represents any one of the above residuals, ε is a white noise, and u can exhibits a serial correlation.

Standard errors appear in parentheses. The sample period is February 1921 through May 1925.

Table 6: Testing the null hypothesis of non-cointegration

		(I) Testing stationarity of the estimated residuals of the cointegrating vector					(II) Testing magnitude and significance of coefficient of the LR residuals in the ECM	
		ADF-test		PH-test		SB-test		
		t_{DF}	Z_{DF}	t_{PP}	Z_{PP}	d_{SB}	φ	t
υ_s		-2.68	-20.03	-0.29	-1.23	0.374 ①	-0.233	-2.760 ⑩
υ_P		-2.89	-20.96	-2.47	-10.02	0.355 ①	-0.219	-3.332 ⑩
υ_{P^*}		-2.72	-15.62	-2.26	-8.05	0.282	-0.047	-1.240
Critical values	1%	-4.45	-35.4	-4.45	-35.4	0.511		-2.62
	5%	-3.75	-27.1	-3.75	-27.1	0.336		-1.95
	10%	-3.36	-23.2	-3.36	-23.2	0.322		-1.61
Table		b.9(3)	b.8(3)	b.9(3)	b.8(3)			b.6(1)

(III) Johansen-test

	Eigenvalue	Likelihood ratio	5% critical value	1% critical value	Hypothesized no. of CIVs
	0.264	25.15	29.68	35.65	None
	0.120	9.81	15.41	20.04	At most 1
	0.066	3.13	3.76	6.65	At most 2
Table			B.11(3)	B.11(3)	

Notes: υ_s is the residuals when the $/£ nominal exchange rate is normalized, υ_P and υ_{P^*} are the residuals when the US-WPI and UK-WPI are normalized, respectively. ADF-test, PH-test and SB-test denote augmented Dickey-Fuller, Phillips-Hansen and Sargant-Bhargava, respectively. φ is the coefficient of the long run error in the ECM. The sample period is February 1921 through May 1925, during which a floating $/£ exchange rate regime prevailed. The estimated statistics outlined by ①, ⑤, and ⑩ indicate significant at 1, 5 and 10 percent levels, respectively. The critical values for t_{DF}, Z_{DF} and d_{SB} are taken from Granger and Yoo (1987). The critical values for t_{PP} and Z_{PP} are from Phillips and Ouliaris (1990). The critical values t are taken from the conservative values of Dickey and Fuller. The critical values for J-test are taken from Johansen (1988) and Johansen and Juselius (1990). For convenience, the last row indicates the appropriate table in Hamilton (1994). The table presents unambiguous evidence in favor of the nonstationarity of the residuals implying inability to reject the null hypothesis of no cointegration. The Johansen test supports this conclusion. However, the coefficients of the long run residuals in the ECM are significant in two cases and bounded between zero and -1, indicating a cointegration.

Table7: The short run dynamic- the Error Correction Model

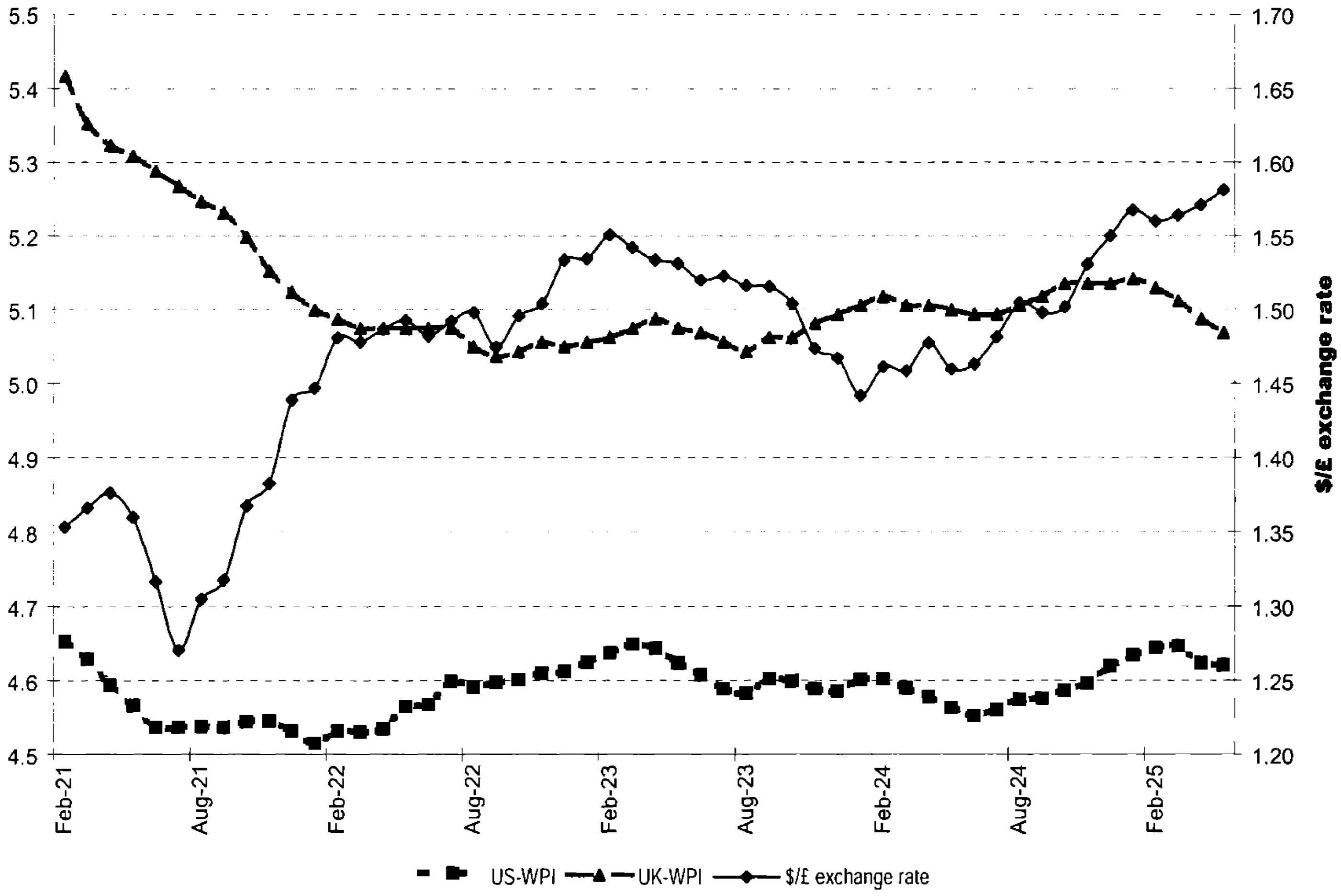
	c_1	φ	ψ_1	ψ_2	ψ_3	ψ_1^1	ψ_2^1	ψ_3^1	ψ_1^2	ψ_2^2	ψ_3^2
Two step Engle-Granger procedure											
s	-	**-0.233**	-	**0.656**	**-0.464**	-	-	-	**0.372**	-	-
	-	(0.084)	-	(0.195)	(0.171)	-	-	-	(0.126)	-	-
p	-	**-0.219**	**0.189**	**0.416**	-	-	**0.279**	-	-	-	-
	-	(0.066)	(0.081)	(0.105)	-	-	(0.113)	-	-	-	-
p*	-	**-0.047**	**-0.157**	**0.478**	-	-	-	-	-	-	-
	-	(0.038)	(0.084)	(0.099)	-	-	-	-	-	-	-

Notes: s is the $/£ nominal exchange rate, p and p* are the WPI of the US and UK, respectively. The estimated regressions are the following parimonious model estimated by OLS:

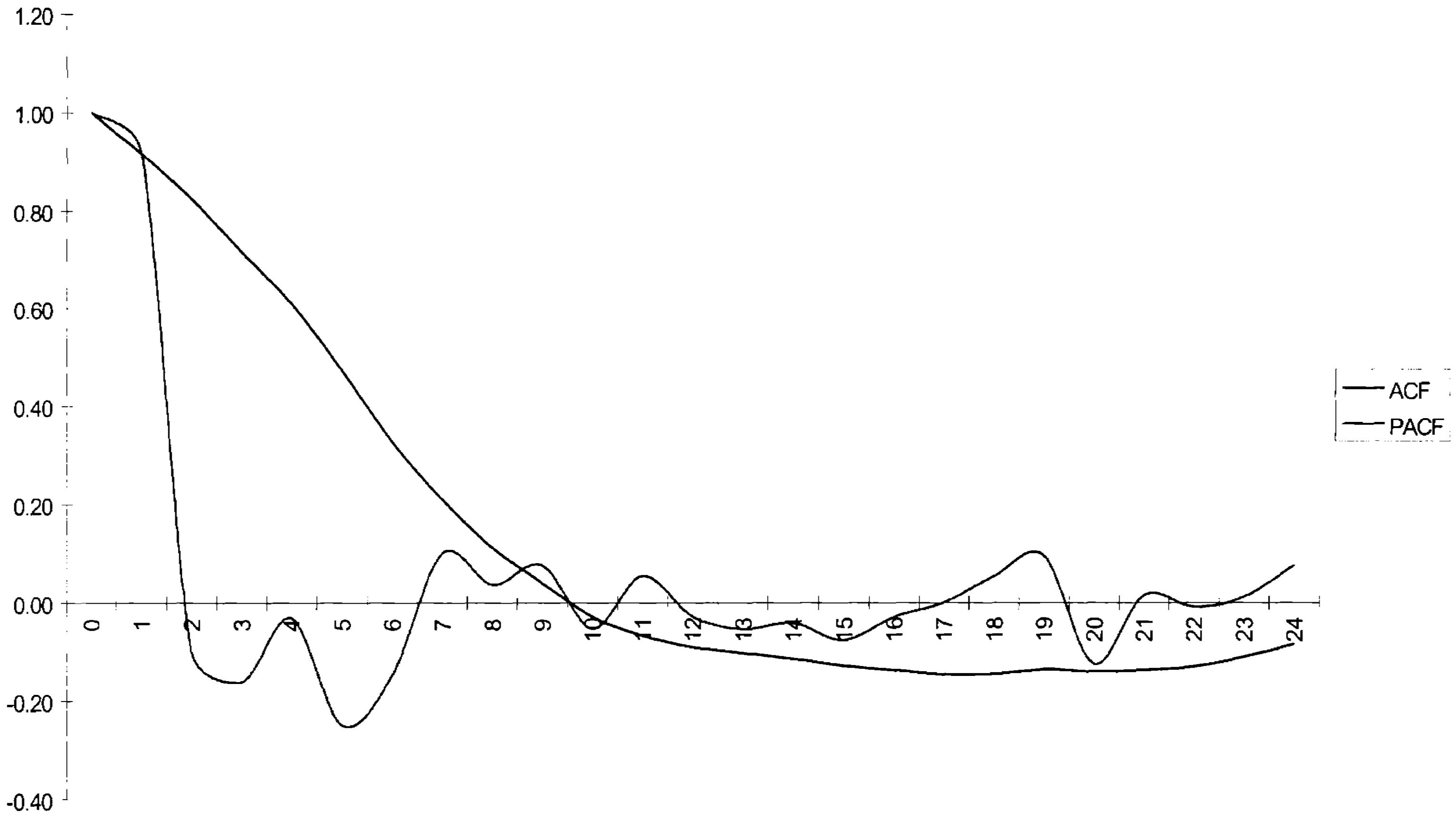
(14) $\psi\Delta x_t = c_{1t} + \varphi \upsilon_{t-1} + \psi^1 \Delta x_{t-1} + \psi^2 \Delta x_{t-2} + \ldots + \psi^h \Delta x_{t-h} + u_t$, where x = (s,p,p*)'.

The sample period is February 1921 through May 1925. The estimates outlined by ①, ⑤, and ⑩ are significant at the 1, 5 and 10 percent levels, respectively. The critical values were taken directly from the appropriate table in Hamilton (1994) as indicated by the last row. For the original source, see in the body of the above test or Hamilton (1994). The one step Johansen procedure was not used since the Johansen test cannot reject the hypothesis of no cointegration.

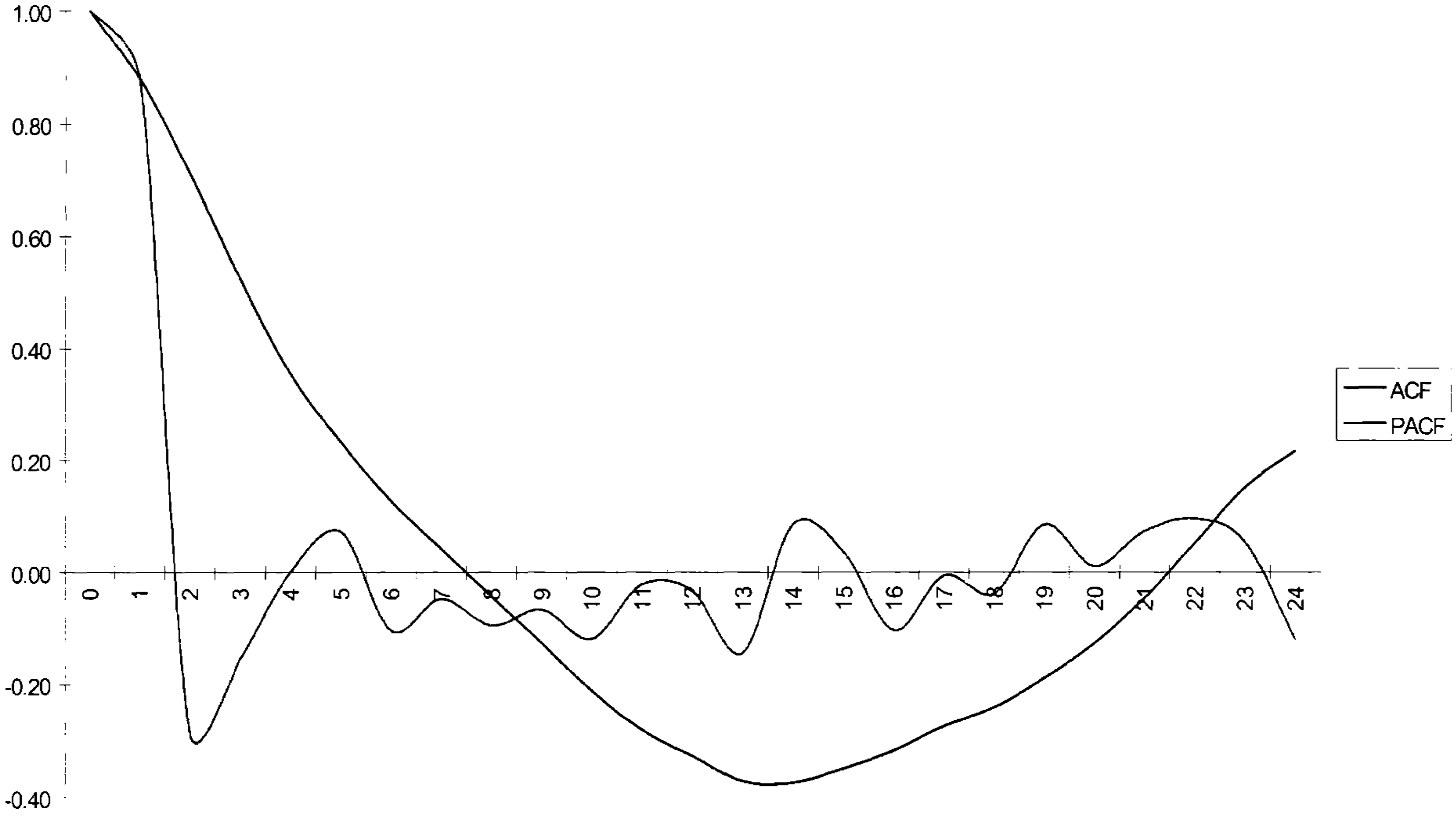
Graph 1: The $/£ spot exchange rate and the US-WPI UK-WPI



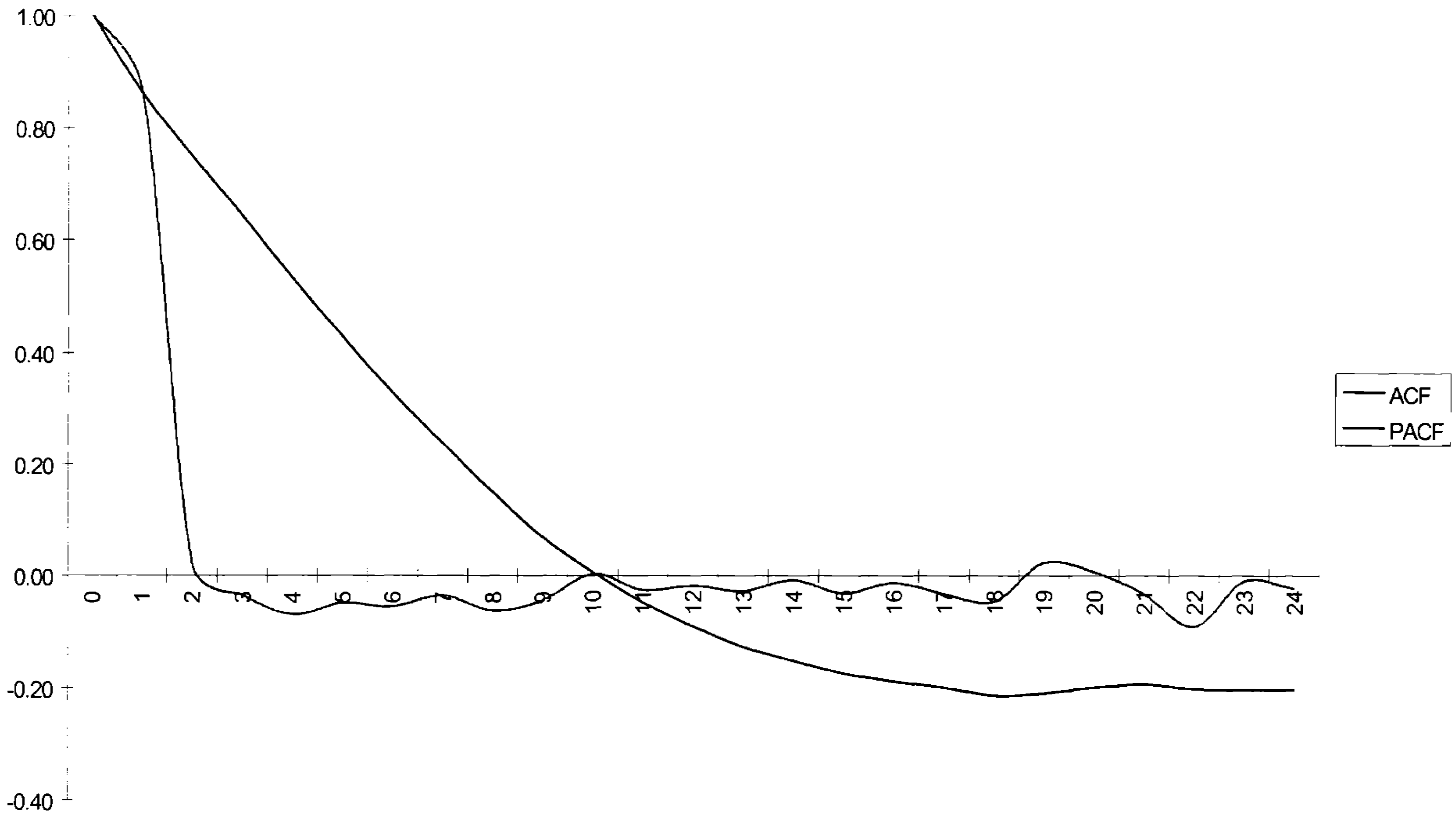
Grapg 2.1: The correlogram and partial correlogram of the $/£ exchange rate



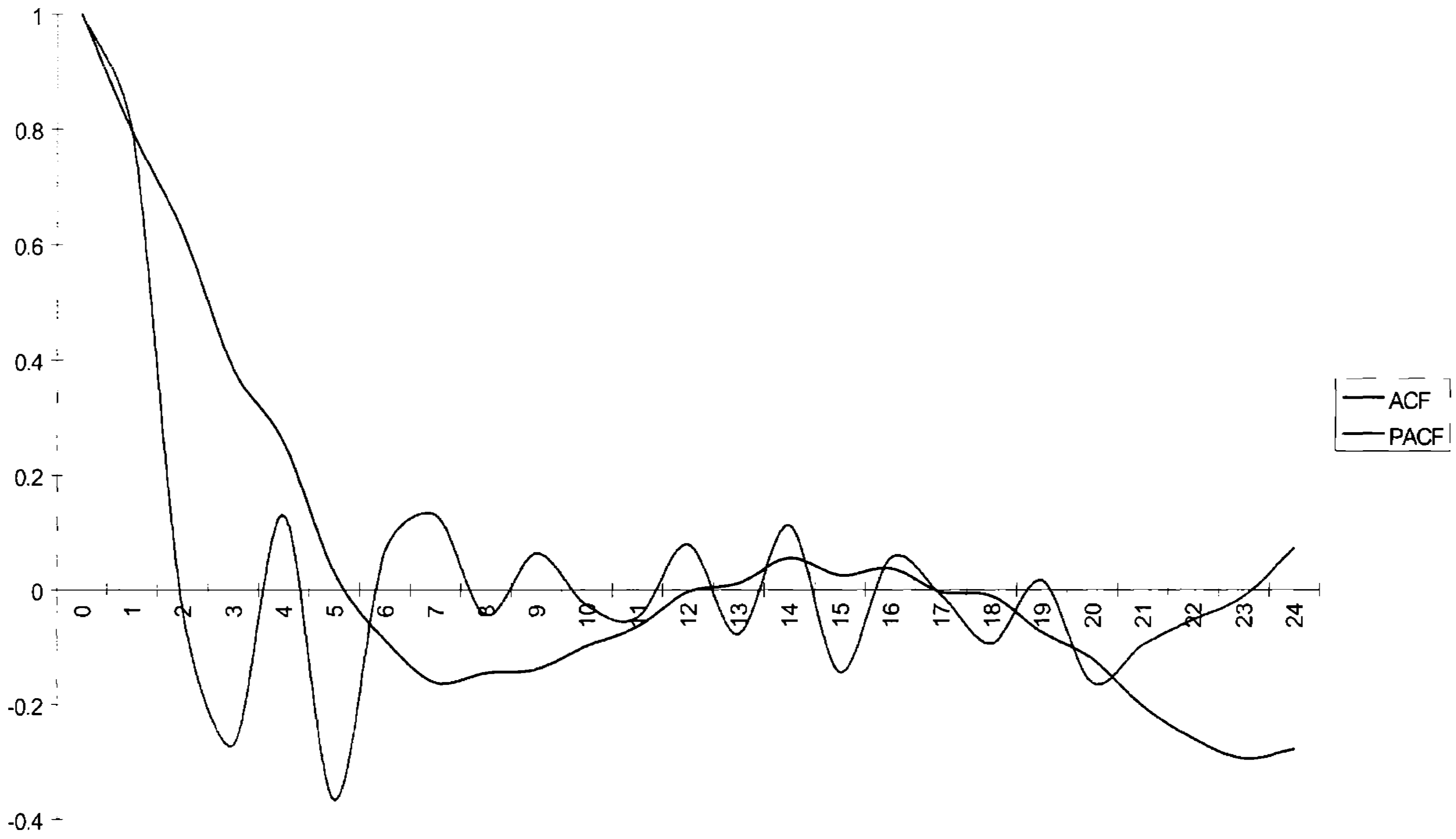
Grapg 2.2: The correlogram and partial correlogram of the US-WPI



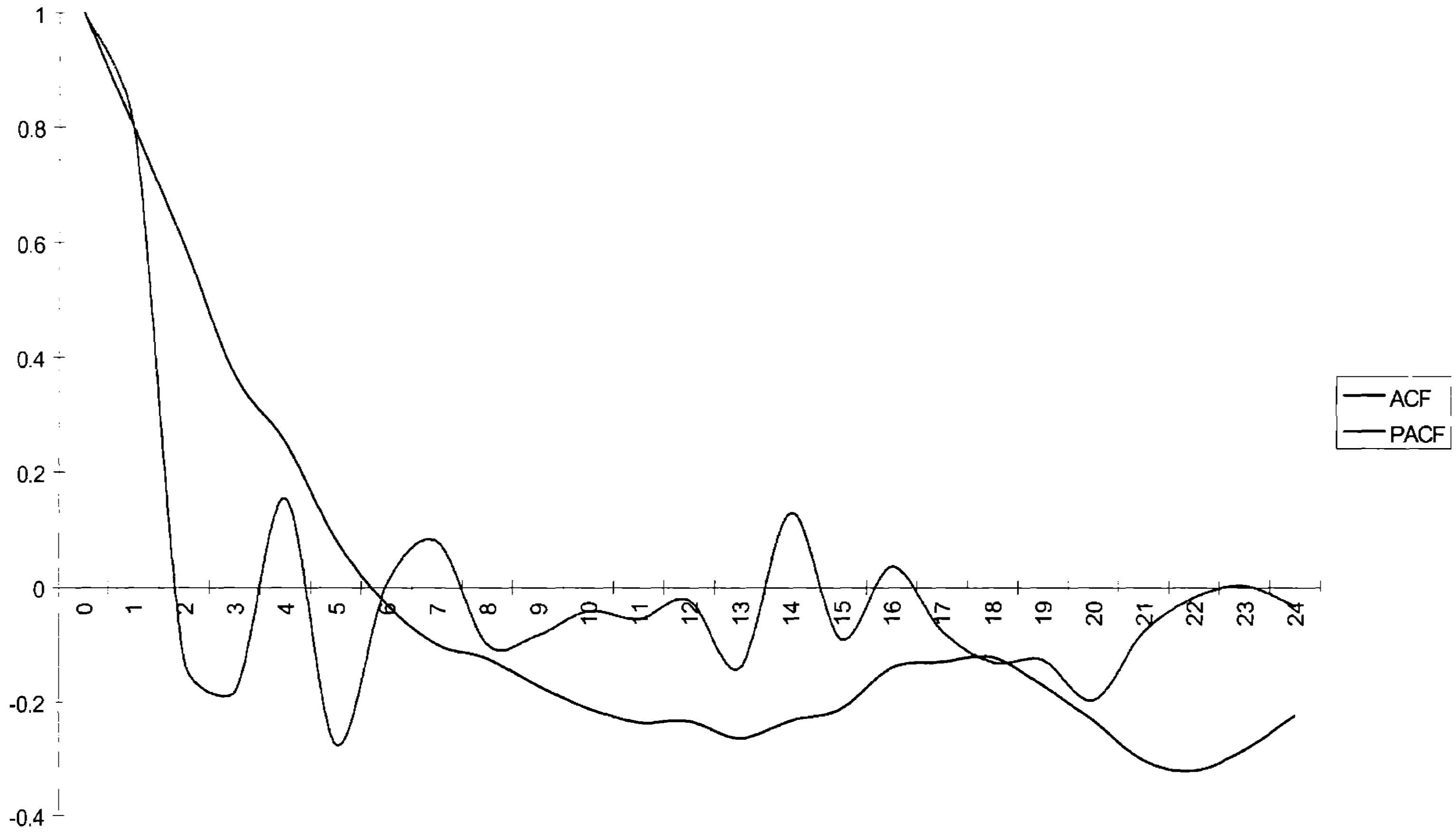
Grapg 2.3: The correlogram and partial correlogram of the UK-WPI



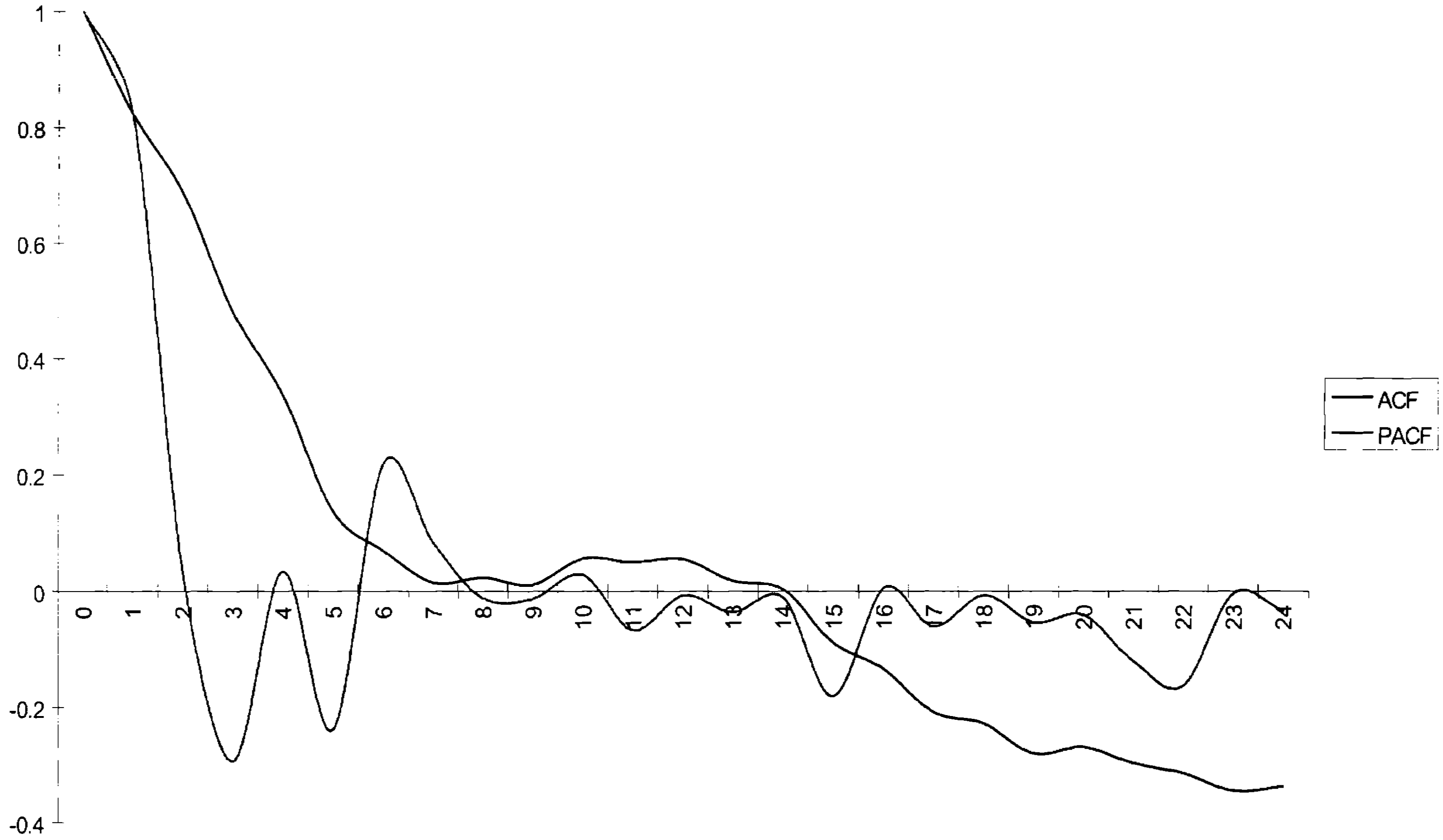
Grapg 3.1: The correlogram and partial correlogram of the long run residuals when $/£ exchange rate is normalised



Grapg 3.2: The correlogram and partial correlogram of the long run residuals when US-WPI is normalised

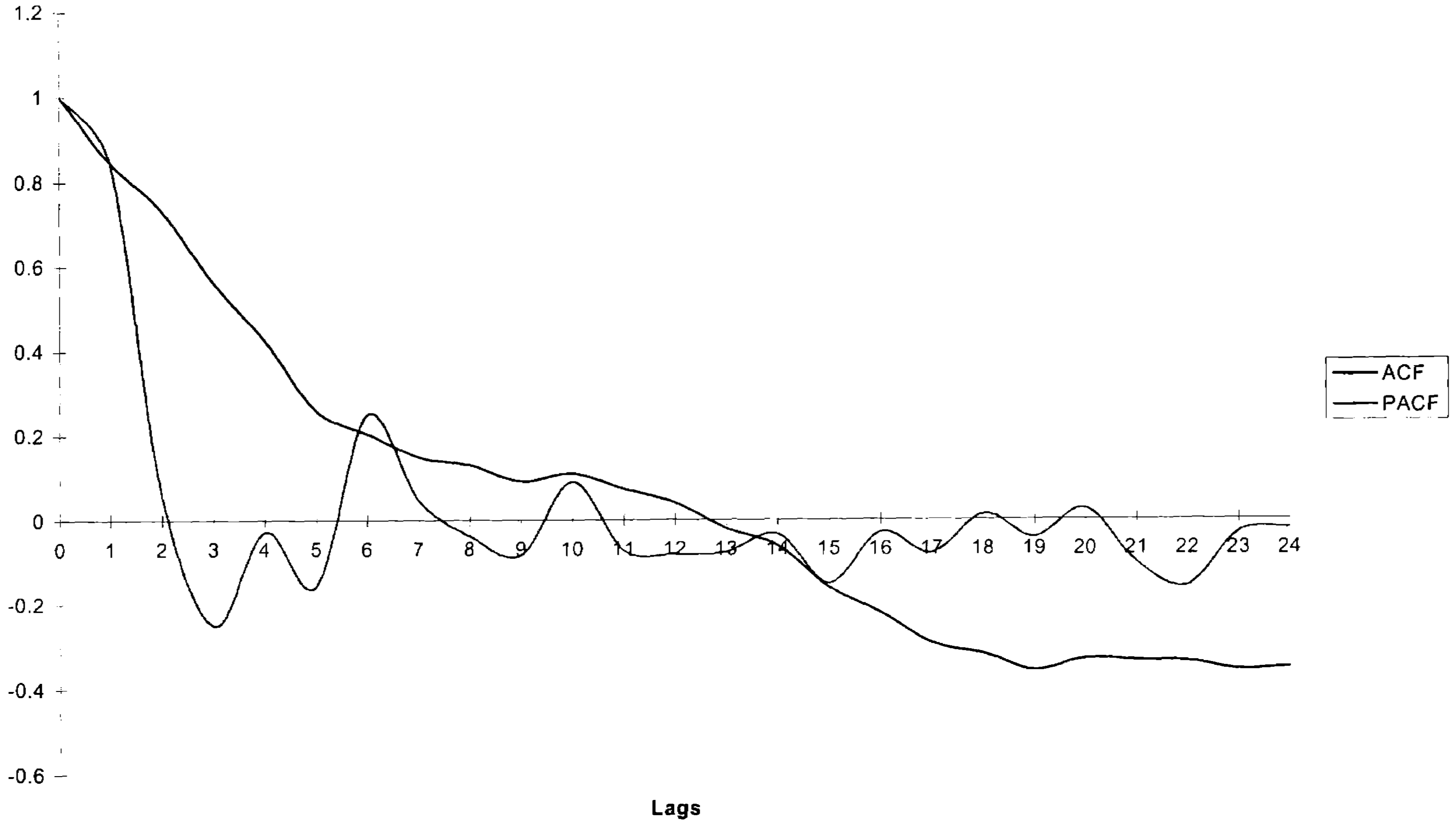


Grapg 3.3: The correlogram and partial correlogram of the long run residuals when UK-WPI is normalised



Graph 4: The $/£ real exchange rate and the PPP

Grapg 5: The correlogram and partial correlogram of the residuals when proportionality is imposed



Appendix 1: Data source and methods[113]

In general, I make every attempt to duplicate the data set used in Clements and Frenkel (1980, 1981), Davutyan and Pippenger (1985), Edison (1989), Frenkel (1976, 1978, 1980, 1986), MacDonald (1985), Phylaktis (1992), and Taylor and McMahon (1988). The data base is made up of 52 monthly observations on each variable for the period February 1921 through May 1925 (see Table A1). During this period, a flexible exchange rate regime was adopted between the UK and US. The weekly spot exchange rate are taken from Einzig (1937, appendix 1, pp. 450-458). The exchange rates quoted are the levels nearest the end of the month. The wholesale price indices for the UK and US are from Tinbergen J. (ed.) (1934), pp. 105-106, column 21 and pp. 210-211, column 28, respectively.

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[113] All the data were subject to logarithmic transformation. See Banerjee *et. al.* (1993), Ch. 1.6, for the advantage of this transformation. See also Granger and Hallman (1991) for analysing general transformation.

Table A1: The data base

Date	S	P	P*
Feb-21	3.8700	104.9	225
Mar-21	3.9200	102.4	211
Apr-21	3.9600	98.9	205
May-21	3.8950	96.2	202
Jun-21	3.7300	93.4	198
Jul-21	3.5625	93.4	194
Aug-21	3.6875	93.5	190
Sep-21	3.7350	93.4	187
Oct-21	3.9250	94.1	181
Nov-21	3.9850	94.2	173
Dec-21	4.2150	92.9	168
Jan-22	4.2500	91.4	164
Feb-22	4.3975	92.9	162
Mar-22	4.3850	92.8	160
Apr-22	4.4225	93.2	160
May-22	4.4500	96.1	160
Jun-22	4.4025	96.3	160
Jul-22	4.4475	99.4	160
Aug-22	4.4725	98.6	156
Sep-22	4.3700	99.3	154
Oct-22	4.4625	99.6	155
Nov-22	4.5000	100.5	157
Dec-22	4.6350	100.7	156
Jan-23	4.6400	102.0	157
Feb-23	4.7150	103.3	158
Mar-23	4.6750	104.5	160
Date Continue ..	S	P	P*
Apr-23	4.6350	103.9	162
May-23	4.6250	101.9	160
Jun-23	4.5725	100.3	159
Jul-23	4.5850	98.4	157
Aug-23	4.5550	97.8	155
Sep-23	4.5525	99.7	158
Oct-23	4.5000	99.4	158
Nov-23	4.3650	98.4	161
Dec-23	4.3375	98.1	163
Jan-24	4.2275	99.6	165
Feb-24	4.3125	99.7	167
Mar-24	4.3000	98.5	165
Apr-24	4.3825	97.3	165
May-24	4.3050	95.9	164
Jun-24	4.3200	94.9	163
Jul-24	4.4000	95.6	163
Aug-24	4.5025	97.0	165
Sep-24	4.4725	97.1	167
Oct-24	4.4900	98.2	170
Nov-24	4.6225	99.1	170
Dec-24	4.7125	101.5	170
Jan-25	4.7950	102.9	171
Feb-25	4.7600	104.0	169
Mar-25	4.7775	104.2	166
Apr-25	4.8125	101.9	162
May-25	4.8600	101.6	159
AVG=	4.3748	98.36	168.6
STD=	0.3153	3.61	15.7

Notes:

s is the dollar/pound exchange rate

p is the WPI of the US

p* is the WPI of the UK

Appendix 2: A summary of the empirical research on PPP

Research	Data frequency	Price index	Exchange rate model	Results	Period
(1)	(2)	(3)	(4)	(5)	(6)
Abuaf and Jorion (1990)	AA EM	CPI CPI	PPP PPP	C C	1900-1972 1/1973-1/1978
Adler and Lehmann (1983)	EM EA	CPI WPI	Martingale Martingale	R R	1/1964-5/1985 1870-1975
Apte *et al.*(1994)	EM	WPI	Martingale	R	1/1973-6/1991
Ardeni and Lubian (1989)	EM EM	WPI WPI	PPP PPP	R R	2/1921-8/23 2/1921-8/1925
Ardeni and Lubian (1991)	EM A	CPI WPI	PPP PPP	R C	1/1957-12/1985 1878-1985
Baillie and Selover (1987)	EM	CPI	PPP	R	4/1973-12/1983
Choudhry *et al.*(1991)	EM EM	CPI + WPI CPI + WPI	PPP PPP	C R	10/1950-4/1962 10/1950-4/1962
Corbae and Ouliaris (1988)	AM	CPI	PPP	R	7/1973-9/1986
Davuytan and Pigginger (1985)	EM	WPI + CPI	PPP	C	1970s+1920s
Edison (1985)	EM	WPI	PPP	R	2/1921-5/1925
Edison (1987)	A	GDP-PD	M-PPP	R	1890-1978
Enders (1988)	EM	WPI	PPP	RC RC	1/1960-4/1971 1/973-11/1986
Enders (1989)	A A	WPI WPI	PPP PPP	C C	1862-1878 1879-1913
Frenkel (1978)	EM	WPI + CPI	PPP	C	2/1921-5/1925
Frenkel (1981b)	EM	WPI + CPI	PPP	R	2/1921-5/1925
Hakkio (1984)	Q Q	WPI + CPI WPI + CPI	Martingale Martingale	R C	1/1921-2/1925 3/1970-4/1982
Kim (1990)	A	WPI	PPP	R	1900-1987
Lothian (1990),	A	WPI	PPP	C	1875-1987
Phylaktis (1992)	EM	WPI	PPP	C	1/1923-12/1925
Roll (1979),					
Rush and Husted (1985)				C	
Sercu (1982).					
Stockman (1980)				R	
Taylor and McMahon (1988)	EM	WPI	PPP	C	1/1921-5/1925

Note: All data are seasonally unadjusted. EM and AM denote end and average of the month, Q and A denote quarterly and annually. WPI and CPI denote the wholesale price index and the consumer price index. C and R denote confirmation and refute of the model. Fi, Fl and M denote fixed, flexible and mixed exchange rate regime, respectively. The abbreviations of the countries' names are given in Appendix 3.

Continue

Research	Fl/ Fix/ Mixed	Adjustment speed	Countries examined	Econometric method
(1)	(7)	(8)	(9)	(10)
Abuaf and Jorion (1990)	Mi Fl	10 10	US; CAN, FR, GER, IT, JAP, NET, SWZ, UK	DF + SURE
Adler and Lehmann (1983)	Mi Mi	43 2	All the 46 countries listed in table 5. US; CAN	F-statistic for AR terms F-statistic for AR terms
Apte *et al.*(1994)	Mi	20	US; ASL, AUS, BEL, CAN, DEN, GER, GRE, ID, IT, JAP, MEX, NET, NOR, SF, SP, SWN, SWZ, UK, VEN	IV + SURE
Ardeni and Lubian (1989)	Fl Fl	4 3	US; FR, GER, UK. US; FR, GER	ADF + ECM ADF + ECM
Ardeni and Lubian (1991)	Mi Mi	7 5	US; CAN, FR, GER, ITA, JAP, UK US; CAN, FR, ITA, UK	VR + ECM VR + ECM
Baillie and Selover (1987)		6	US; CAN, FR, GER, JAP, UK	ADF + ECM
Choudhry *et al.*(1991)	Fl Fl	3 2	US; CAN, UK US; CAN	ADF + CI ADF + CI
Corbae and Ouliaris (1988)	Fl	7	US; CAN, FR, GER, IT, JAP, UK	ADF + PP + IC
Davuytan and Pigginger (1985)		4	US; UK	OLS
Edison (1985)		3	US; FR, UK	GIVE + ECM
Edison (1987)		2	US; UK	GIVE + ECM
Enders (1988)	Fi Fl	3	US; CAN, GER, JAP	ARIMA + ECM
Enders (1989)	Fl Fi	2	US; UK	ADF + ECM
Frenkel (1978)		10	US, UK, FR, GER	
Frenkel (1981b)		4	US, UK, FR, GER	
Hakkio (1984)				
Kim (1990)		3		
Lothian (1990)		4	JAP; FRA, UK, US	

Research	Fl/ Fix/ Mixed	Adjustment speed	Countries examined	Econometric method
Phylaktis (1992)		4	GRE; FRA, UK, US	ADF + ECM
Roll (1979),				
Rush and Husted (1985)				
Sercu (1982).				
Stockman (1980)				
Taylor and McMahon (1988)		4	US, UK, FR, GER	ADF + ECM

Appendix 3: Abbreviation of countries' names

#	Abbreviation		Full country name	#	Abbreviation		Full country name
1	ARG	=	Argentina	24	KUW	=	Kuwait
2	ASL	=	Australia	25	LUX	=	Luxembourg
3	AUS	=	Austria	26	MAL	=	Malaysia
4	BEL	=	Belgium	27	MEX	=	Mexico
5	BOL	=	Bolivia	28	NET	=	Netherlands
6	BRA	=	Brazil	29	NIG	=	Nigeria
7	CAN	=	Canada	30	NOR	=	Norway
8	CHI	=	Chile	31	PAK	=	Pakistan
9	CON	=	Congo	32	PAR	=	Paraguay
10	DEN	=	Denmark	33	PER	=	Peru
11	ECU	=	Ecuador	34	PHI	=	Philippines
12	EGY	=	Egypt	35	SIN	=	Singapore
13	FIN	=	Finland	36	SAF	=	South Africa
14	FRA	=	France	37	SPA	=	Spain
15	GER	=	Germany	38	SWE	=	Sweden
16	GRE	=	Greece	39	SWI	=	Switzerland
17	IND	=	India	40	THA	=	Thailand
18	IDO	=	Indonesia	41	UK	=	United Kingdom
19	IRA	=	Iran	42	US	=	United States
20	ISR	=	Israel	43	URU	=	Uruguay
21	ITA	=	Italy	44	VEN	=	Venezuela
22	JAP	=	Japan	45	ZAI	=	Zaire
23	KEN	=	Kenya	46	ZAM	=	Zambia

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