Exchange Rate Exposure Revisited

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Abstract

This paper examines the relation between exchange rate changes and ¯rm value. We introduce a new criterion (\net exports-to-sales ratio") to select ¯rms which are most likely to be a®ected by exchange rate changes. In contrast to previous studies, our criterion captures the counteracting exposures to exchange rate movements arising from the import and export activities of the ¯rm. We ¯nd no evidence of a contemporaneous exchange rate exposure for the industry with the largest net exports-to-sales ratio at the 4-digit SIC level; however, a lagged change in exchange rates signi¯cantly reduces industry returns over a long period (1976-1990). Controlling for size, mature ¯rms from the largest net exporting industry are less exposed than relatively less mature ¯rms. This is consistent with the hypothesis that the ability to hedge exchange rate exposure may be enhanced over time. In contrast to the largely insigni¯cant exchange rate exposure at short horizons, at long horizons the FX exposure is both statistically and economically very signi¯cant.

1 Introduction

All ¯rms operating in the global environment are, in principle, exposed to foreign exchange risk (FX risk). Even in the extreme case where the ¯rm is not engaged in international trade, does not have operations abroad, and has only domestic competitors, it may nevertheless be exposed to FX risk if, for example, its suppliers are exposed to FX risk via international trade, foreign operations or import competition. In light of this, it is surprising that e®orts to detect the e®ects of such exposure (Jorion (1990), Amihud (1993) and Bartov and Bodnar (1994)) have been unable to document signi¯cant evidence of a contemporaneous correlation between exchange rates and ¯rms' values.[1]

The purpose of this paper is to examine the impact of exchange rate movements on ¯rms' values while correcting for possible problems that have arisen in past studies. Previous inability to document FX risk e®ects may be attributed to either or both of the following explanations: a) Previous research has concentrated on the exporting character of the ¯rms in the sample but has not taken into account their possible importing activities, which naturally hedge FX exposure arising from the exporting activities; and b) Previous tests based on short-horizon regressions (using monthly or quarterly data) may not capture the e®ects of the long swings of the dollar and lack statistical power.[2] [3]

[1] In particular, Jorion (1990) ¯nds that only 15 out of 287 U.S. multinationals in his sample are signi¯cantly exposed to FX risk, which is slightly higher than the 5 percent expected to be obtained by chance; Amihud (1993) ¯nds no signi¯cant exposure even for the portfolio comprised of the 8 largest exporting companies where, on average, exports account for 24 percent of their total sales.

[2] Feldstein (1988) notes: \the dollar has experienced three big swings: The ¯rst of these is marked by a sustained rise of foreign currencies against the dollar; between the beginning of 1977 and the end of 1979, the mark gained 3.3 percent against the dollar, the franc gained 21 percent and the pound 26 percent. This was followed by a ¯ve year surge in the dollar at the end of which, these 3 European currencies fell 60 to 90 percent (in log terms) against the dollar. Early in 1985, foreign currencies once more began to rise, gaining 50 to 70 percent against the dollar by the end of 1987" (see Engel and Hamilton (1990) for a formal model of long swings).

[3] A third possible explanation is that ¯rms use foreign currency derivatives to reduce or eliminate their

More speci¯cally, to examine a), ¡ and given the absence of ¯rm-level information on imports and exports ¡, we design a sample selection process to sample ¯rms which belong to the industry, at the 4-digit SIC level, with the largest \net exports-to-sales ratio"; and to examine b), we perform long-horizon regressions using overlapping returns of the largest net exporting industry on an exchange rate index. In addition, we also perform long-horizon regressions using individual ¯rm returns from the automotive industry. The automotive industry lends itself to several interesting tests of the e®ects of FX movements on ¯rm value. For example, i) is the adverse impact on pro¯tability of the U.S. automotive ¯rms, caused by the appreciation of the dollar in the ¯rst half of the 1980s, re°ected in their FX exposure? and ii) given the very competitive environment of the automotive industry and the rivalry with the Japanese automotive ¯rms, how has the value of the U.S. ¯rms been a®ected by the dollar/yen movements in the short and the long run?

Finally, we examine evidence documented by Bartov and Bodnar (1994) that a lagged change in the exchange rates signi¯cantly a®ects stock returns. This raises the possibility that the lack of evidence of a contemporaneous exposure is a special form of market ine±ciency, namely, lagged adjustment. Bartov & Bodnar (1994) use an updating procedure to select each year, a di®erent sample of ¯rms (based on their accounting exposures), suggesting that mispricing can occur in the short term. We examine whether this form of mispricing can also be present for a long period of time for the industry with the largest net exports-to-sales ratio. In contrast to Bardov and Bodnar's approach, where the nature of the ¯rms' operations in their sample is not clearly identi¯ed, an advantage of our approach is that it re°ects the importance of a ¯rms' real operations (imports and exports) in their exchange rate exposure.

We ¯nd no signi¯cant contemporaneous exposure for the industry with the largest net exports-to-sales ratio. However, in line with Bartov and Bodnar (1994), we ¯nd evidence

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exposures (see e.g., Stulz (1984), Smith and Stulz (1985) and Froot, Scharfstein and Stein (1993) on the optimality of hedging). Although this explanation is not directly investigated here, it is clear that the use of derivatives by ¯rms in a risk-reduction manner should bias our tests against ¯nding any exposure.

of mispricing, namely, that a lagged change in the exchange rates significantly affects the industry returns. In contrast with Bartov and Bodnar though, this mispricing is present for the entire period of our tests (January 1976 - December 1990). The existence of lagged exposure in our framework is consistent with the lagged release of imports and exports data by the U.S. department of Commerce. In addition, the positive sign of the lagged exposure coefficient is consistent with our hypothesis, indicating that for the industry with the largest net exports-to-sales ratio, an appreciation of the dollar reduces its returns. Mature firms in this industry (firms which have existed since 1976) are less exposed to FX risk than less mature firms (firms which have existed since 1982). This may be evidence of a learning behavior on the part of the firm, in that firms learn over time how to better manage FX risk. These results are robust to the use of alternative exchange rate indices and econometric specifications in the estimation of the exchange rate exposure.

At long horizons, firms' exchange rate exposures differ markedly from their FX exposures at short horizons. We find strong evidence that firms are more significantly exposed to exchange rate changes at long horizons than at short horizons. For the largest net exporting industry, the sign of the long-horizon exposure is consistent with our hypothesis that an appreciation of the dollar significantly reduces the industry returns at long horizons. For the U.S. automotive firms, the signs of the long-horizon exposures are consistent with reports on their overall competitive strategies (e.g., imports, exports, cost restructuring and foreign acquisitions).

The paper has implications for corporate exchange risk management policies: a) Imports may be effectively used to protect firms from FX exposure arising from exports and, b) The long-horizon effects of the exchange rate movements may be significant for corporations, even though their short-term exposures are small.

The rest of the paper is organized as follows: section 2 motivates and formulates the hypotheses and presents the models and variable definitions; section 3 describes the sample selection procedures and the data that we use in our tests; section 4 presents the tests and results; and section 5 concludes.

2 Hypotheses Formulation and Models Used

Dumas (1978), Adler and Dumas (1984) and Hodder (1982) de¯ne economic exposure to exchange rate movement as the regression coe±cient of the real value of the ¯rm on the exchange rate, across states of nature. The de¯nition does not imply causality, namely, that exchange rate changes cause changes in ¯rms' values, or vice versa. Indeed, in Adler and Dumas (1980), stock prices and exchange rates are both endogenous variables and are determined simultaneously. However, for an individual ¯rm (industry), it can be safely assumed that exchange rates are exogeneous.

An unexpected appreciation of the dollar negatively a®ects a U.S. exporting ¯rm with revenues in foreign currencies, regardless of the possible adjustment in foreign currency export prices.[4] This is true as long as the ¯rm's cost structure is not a®ected by exchange rate movements. If the ¯rm is using imported, or simply, internationally priced inputs, then its cost structure will be positively a®ected by an unexpected appreciation of the dollar (i.e., importing becomes cheaper) and the overall FX exposure of this exporting ¯rm will be reduced, or even reversed. In general, one would expect that for a given ¯rm, the higher the export-to-sales ratio and the lower the import-to-sales ratio, the higher its overall exposure to exchange rate risk. Hence, the industry with the largest \net exports-to-sales" ratio should be the most signi¯cantly exposed industry to exchange rate movements. For this industry, an appreciation of the dollar should negatively a®ect its returns.

Formally, the ¯rst hypothesis (**H1**) that we test is whether, \an appreciation of the real exchange rate of the U.S. dollar against the foreign currencies negatively a®ects U.S. ¯rms with the largest net exports-to-sales ratio."

[4]See, for example, Giovannini (1985) and Krugman (1987), Levi (1993), Shapiro (1975), .

We test **H1** using the following model for each firm/portfolio return i:

$$R_{it} = \beta_{0i} + \beta_{1i} R_{mt} + \beta_{2i} FXI_t + \varepsilon_{it}; \; t = 1; \ldots; T \tag{1}$$

where,

R_{it} is the real rate of return on the i firm's (portfolio of firms) common stock in period t;

R_{mt} is the real rate of return on the market portfolio in period t;

FXI_t is the real rate of return on a moving, trade-weighted exchange rate index, measured as the exchange rate of the U.S. dollar against the foreign currencies.

Model (1) is in the spirit of Adler and Dumas (1984) definition of exchange rate exposure and β_{2i} is the exchange rate exposure coefficient (similar models were used in Jorion (1990) and Amihud (1993)). This specification assumes that exchange rates and stock returns follow a random walk process and hence the rate of return captures the unanticipated movements. However, we also estimate a model, where we account for possible autocorrelation in the exchange rate index. We should note that there is little difference between nominal and real exposure in our framework, since the largest percentage of variation comes from exchange rates and not from inflation. Similarly, there is little difference between using excess returns (returns over the risk-free rate) and simple returns, since the variation in interest rates is also relatively small compared to the variation in exchange rates.[5] To test restrictions across equations (i.e., the joint hypothesis that firm exposure coefficients within an industry are all zero), we employ seemingly unrelated regressions (SUR) using the following model:

$$R_{it} = \beta_{0i} + \beta_{1i} R_{mt} + \beta_{2i} FXI_t + \varepsilon_{it}; \;\; i = 1; \ldots; 5; \; t = 1; \ldots; T \tag{2}$$

[5] For example, over the period 1971-1987, the annualized volatility of the dollar/mark exchange rate change was 12% versus a volatility of 3% for the U.S Treasury bill rate and 1.3% for the U.S inflation.

where the variables are as defined in model (1).

Bartov and Bodnar (1994) employed a sample of U.S. firms that have consistently reported large accounting exposures (foreign-currency adjustments) on their past annual financial statements. In addition, these exposures were negatively correlated with the corresponding changes in the dollar. For this sample, they find no significant evidence of a contemporaneous correlation between firms' values and exchange rate changes.[6] However, they do find that a lagged change in the exchange rates is significantly correlated with equity returns. Bartov and Bodnar explore further the possibility that this might be due to mispricing. The authors use an updating procedure to select a different sample of firms each year based on information about their accounting exposure for the past five years. This suggests that mispricing can occur in the short term. In this paper, we examine whether for the largest net exporting industry, systematic mispricing occurs over a long period of time. Given that information on imports and exports is released with a lag of 45 days by the Department of Commerce, it is conceivable that the financial markets react to the news with a lag.

Formally, the second hypothesis (**H2**) that we test is whether, "a lagged appreciation of the real exchange rate of the U.S. dollar against the foreign currencies negatively affects U.S. firms with the largest net-export-to-sales ratio"

We test **H2** using the following model:

$$R_{it} = \alpha_i + \beta_i R_{mt} + \delta_{1i} FXI_t + \delta_{2i} FXI_{t-1} + \epsilon_{it}; \ t = 1, \ldots, T \qquad (3)$$

where R_{it}, R_{mt} and FXI_t are as defined in model (1) and FXI_{t-1} is defined as the lagged change in the exchange rate index. In the above model, δ_{2i} represents the lagged exposure.

Finally, we examine whether the inability to capture FX exposure was due to the use of short-horizon regressions. There are both economic and statistical reasons to support the use

[6] Recently, Palia and Thomas (1996) argue that there is significant economic exposures for firms with large accounting exposures that were positively correlated with the corresponding changes in the dollar.

of long-horizon regressions to capture FX exposure. The dollar has exhibited a behavior of long swings during the period that we examine.[7] Therefore, we would expect stock returns to exhibit a signi¯cant long-term exposure to the dollar's movements. The use of short-horizon regressions may segment the expected long-term FX exposure rendering the long-term trend undetected. Also, the lack of a perfect long-term ¯nancial hedge may leave ¯rms exposed to long-term exchange rate movements. As Mello and Parsons (1995) argue, while short-term exposure could be perfectly hedged through the available short term instruments (i.e., futures), the same is not true when hedging long-term exposure.

In addition, long-horizon regressions have been shown to increase the statistical power of the tests in a number of cases (Campbell (1993)).[8] Furthermore, it has been argued that the use of overlapping observations in the calculation of long-horizon relations provides more e±cient estimators than the use of nonoverlapping observations (Hansen and Hodrick (1980)). Boudoukh and Richardson (1993) show that there are e±ciency gains of long-horizon regressions when using overlapping observations as long as the autocorrelation of the regressor is low. Also, Stambaugh (1993) suggests that violations of OLS standard assumptions, such as heteroscedasticity, can make long-horizon regressions more e®ective. In this paper, we estimate the long-horizon (3, 6, 9 and 12 month) exposures for the industry with the largest \net exports-to sales ratio" and for the sample of the U.S. automotive ¯rms and compare them to their short-horizon (1 month) exposures.

[7] Engel and Hamilton (1990) develop a statistical model of exchange rate dynamics as a sequence of stochastic, segmented time trends. Their tests reject the null that exchange rates follow a random walk, in favor of their model of long swings.

[8] Many authors have examined the relationship between long-horizon stock returns and a variety of instrumental variables, such as, contemporaneous dividend yields and in°ation (e.g., Fama and French (1988), Bekaert and Hodrick (1992) and Boudoukh and Richardson(1993)). In most of these studies, the statistical power as measured by the adjusted R^2 signi¯cantly increased with an increase in the horizon. Although Goetzman and Jorion (1993) suggest that there may be biases involved in the above studies, these biases are due to the fact that \the right-hand side variables are correlated with lagged dependent variables, instead of being predetermined as assumed in standard statistical models". This is clearly not the case in our tests, where exchange rates can be assumed exogenous.

Formally, the third hypothesis (**H3**) that we test is whether "a lagged appreciation of the real exchange rate of the U.S. dollar against the foreign currencies negatively affects the U.S. industry with the largest net exports-to-sales ratio at long horizons."

For the automotive industry, we examine whether the signs of the long-horizon exposures are in line with the overall strategies of the automotive firms in different subperiods between January 1976 and December 1990.

We test **H3** using the following model:

$$\sum_{i=1}^{J} R_{t+i} = \alpha_J + \beta_J \sum_{i=1}^{J} R_{mt+i} + \gamma_J \sum_{i=1}^{J} FXI_{t+i} + \epsilon_t(J); \; t = 1, \ldots, T \qquad (4)$$

where, R_t, R_{mt} and FXI_t are defined as in model (1). The difference in model (4) is that this model uses overlapping data for R_t, R_{mt} and FXI_t and that the returns are calculated over a long-horizon according to J. In the above model, γ_J represents the long-horizon exposure.

3 Sample Selection and Data

3.1 The Index

For the purposes of this paper we use a real, trade-weighted, monthly dollar index (RX-101) compiled by the Federal Reserve Bank of Dallas.[9] This index differs from those used in earlier studies in two ways: i) by the method used to construct trade weights and ii) by the selection of currencies against which to measure the dollar. In particular, moving trade

[9] To be consistent with Adler and Dumas (1984) definition of exchange rate exposure, the original exchange rate index was inverted so that the index that we ultimately used in our tests is measured in US dollars per unit of foreign currencies.

weights are employed rather than weights that are tied to certain years or trading °ows.[10] Also, the RX-101 index utilizes 101 U.S. trading partners, in contrast to only 15 used to construct the Morgan Guaranty Trust Company of NY (MG) exchange rate index (used by Amihud (1993)) and 22 for the IMF's MERMA (used by Jorion (1990)). In addition, the Morgan Guaranty Trust indices include Switzerland among the 15 countries, which ranks only as the U.S.'s 20th trading partner, and do not include Mexico, which is U.S.'s 3rd largest trading partner, nor countries from the Western hemisphere, which together accounted for more than 37% of total U.S. trade in 1985. However, to check the robustness of our results to alternative exchange rate indices, we also use the MG index to test **H1** and **H2**.

To correctly estimate FX exposure, it is important to examine possible autocorrelation in the indices that we use. Individual exchange rates have been shown to follow a nearly random walk process (see e.g., Mussa (1979), Meese and Rogo® (1983)), which implies that the actual changes in the exchange rates represent the unexpected changes. However, to the extent that our indices are autocorrelated, this is no longer true. In this case, as pointed out earlier, the residuals from the regression of the exchange rate index on its lagged change should be used. The autocorrelations for the indices that we used are generally small. In particular, the autocorrelation is 0.22 for the RX-101 and 0.18 for the MG. In the next section, we have included tests that correct for the presence of autocorrelation in the exchange rate indices. Finally, in the tests of **H3**, in addition to the RX-101, we also use the real Dollar/Yen exchange rate, supplied by the Federal Reserve Bank of Dallas.

3.2 The \net exports-to-sales" ratio

We use the following procedure to select the sample that we use in our tests: At the 2-digit SIC level, we calculate the di®erences between exports to sales and imports to new

[10] Moving weights could potentially present a problem in our tests; if participants in ¯nancial markets do not ex-ante know these weights, they will not know to what FX changes to respond to. This potential problem is alleviated when we examine the exposures to the dollar/yen and to alternative indexes.

supply, according to statistics given by the 1990 report of Industry and Trade Statistics of the U.S. Department of Commerce, and rank them from highest to lowest. We then select the industry with the highest net exports at the 2-digit SIC code (SIC code 35 : Industrial Machinery and Equipment) and repeat the above procedure within the industry at the 4-digit SIC level. The industry that ranks ¯rst (among the industries at the 4-digit SIC level for which a substantial number of ¯rms with no missing data between 1976 and 1990 exists) was subsequently used in our tests (Electronic Computers, SIC code 3571). The Electronic Computers industry exports approximately 37.4% of its production.

We obtained data for all ¯rms in the Electronic Computers industry for which monthly returns were available in the CRSP database for the period January 1976 to December 1990. The ¯ve ¯rms for which return data exist for the entire time period are Tandy, Commodore International, Alpha, Qantel and Electronic Associates. We use both, an equally weighted portfolio of individual ¯rms' returns and a value weighted industry portfolio, which was constructed with weights proportional to the percentage of each ¯rm's total assets within the industry. Data for electronic computer ¯rms for which monthly returns exist only from 1982 onwards are also obtained (sample of less mature ¯rms).

3.3 Other Data

In addition to using the Electronic Computer industry to test for long-term exposure, we also use all the ¯rms that belong to the Motor Vehicles and Car Bodies industry (SIC code 3711) and have complete return data for the period between January 1976 and December 1990. A total of ¯ve companies are represented: GM, Ford, Chrysler and two smaller in size companies, Navistar International, which specializes in trucks and Federal Mogul which engages mainly in automotive parts. To adjust the nominal stock returns for in°ation we use the in°ation index PUNEW (CPI-U) retrieved from CITIBASE. The CRSP monthly value weighted market index was used as the market index in all tests.

4 Tests and Results

4.1 Control for the Importing Activities of Exporting Firms; Test of H1.

We test **H1** for the entire period, January 1976 to December 1990, and for three equally divided subperiods, (1976-1980), (1981-1985) and (1986-1990). Given that our index is expressed in U.S. dollars per unit of foreign currency, it decreases with an appreciation of the dollar. According to our hypothesis, the value of the net exporting ¯rms should also decrease with an appreciation of the dollar; hence, we should expect a positive exposure coe±cient. For the entire period, as well as for all three subperiods, the equally weighted portfolio returns of the Electronic Computers industry (SIC code 3571) exhibited no signi¯cant exposure to the changes of the real exchange rate index RX-101. In addition, the signs of the FX exposure coe±cients (¯$_2$) are negative in all periods (see table 1). This is evidence against our hypothesis that an appreciation of the dollar would reduce the Electronic Computers industry returns.[11]

To examine individual ¯rm exposure within the electronic computer industry and to test the hypothesis that all FX exposures across ¯rms are jointly zero, we employ a SUR model (model 2) with as many equations (5) as the number of ¯rms in the industry with available data. The results are shown in table 2; the ¯rms listed are in descending order according to size, as measured by the total value of assets at the end of year 1990. For the entire period, January 1976 to December 1990, only Commodore exhibits marginal signi¯cant exposure (for a one-tailed test at the 10 % level) and a sign (positive) consistent with our alternative hypothesis. This is evidence that during the period January 1976 to December 1990, an unexpected appreciation of the dollar reduced Commodore's stock return. In contrast, Qantel exhibits signi¯cant exposure (for a one-tailed test at the 10 % level) but the sign (negative)

[11] Correcting for heteroscedasticity and serial correlation in the error structure does not materially a®ect our results.

is the opposite from what we expected. For Qantel, an unexpected appreciation of the dollar actually increased its stock return. None of the remaining firms were significantly exposed to FX movements during the entire period (1976-1990).

Subperiod results are equally puzzling since none of the firms are significantly exposed to FX risk, except for the subperiod 1985-1987, where three out of the five firms in the industry are significantly exposed to FX risk; however, the signs of two of these are again the opposite from what was hypothesized. A test examining whether all exposure coefficients are jointly zero is shown in the bottom panel of table 2. Only in the last subperiod (1988-1990), can we reject this hypothesis at standard significance levels. In conclusion, the above tests show little evidence of significant contemporaneous exposure for the largest net exporting firms.

4.2 Test for the Existence of Mispricing; Test of H2.

To test **H2**, we use the sample of firms from the Electronic Computers industry (highest net exports to sales ratio) and employ model (3). Before using model (3) in practice, we test whether the model is well specified. We start by defining a model of rationally distributed lags, where the number of lags is known ex-ante. We consider 24 lags for our model and apply the Akaike's information criterion to choose between the alternatives.[12] For our sample, a model with one lag was chosen according to this criterion.

The results are surprising, but in line with what Bartov and Bodnar (1994) find. As shown in table 3, there appears to be mispricing, as a lagged change in the exchange rate significantly affects the portfolio returns of the firms in that industry. Our results differ however from Bartov and Bodnar's in one significant aspect: Bartov and Bodnar (1994) use

[12] Akaike's information criterion assumes the form

$$AIC(n) = \ln(\sigma_n)^2 + 2n/T \qquad (5)$$

where σ_n is the MLE evaluated under the assumption that $n = n^*$ (24 in our case) and an estimate $n_{est}(AIC)$ of n^* is chosen, so that AIC assumes its minimum for $n = n_{est}$.

quarterly data and hence their ¯ndings indicate that a one quarter lagged exchange rate change a®ects returns. Instead our use of monthly data indicates that a one month lagged exchange rate change a®ects returns.

In the ¯rst panel of table 3 we report results for the equally weighted (value weighted) industry portfolio of the more mature ¯rms IPMW (IPMVW) for which data exist for the period between 1976 and 1990, and two size portfolios, IPML (large, mature ¯rms) and IPMS (small, mature ¯rms). Both the equally weighted industry portfolio returns and the portfolio of the larger ¯rms, IPML, are signi¯cantly a®ected by a lagged change in the exchange rate index for the entire period. A 1% lagged appreciation of the dollar decreases the return on the equally weighted industry portfolio by 1.09%. In contrast, the portfolio of the small ¯rms is not signi¯cantly exposed to a lagged change in exchange rate. This ¯nding that the larger ¯rms in the industry exhibit stronger FX exposure can be reconciled with Smith and Stulz (1985) whose theoretical model predicts that smaller ¯rms should hedge more than larger ¯rms. In all the various portfolios that we examine, the signs of the lagged exposures (positive) are consistent with our alternative hypothesis, namely that an appreciation of the real exchange rates at time t ¡ 1 will reduce the industry returns at time t. Bartov and Bodnar (1994) provide an intuitive explanation of why a lagged exposure to exchange rates might exist, based on the timing of the release of relevant information to the ¯nancial markets. One possible explanation is that our results re°ect the fact that information on industry imports and exports is released to the public by the U.S. Department of Commerce with a lag of 45 days. The ¯nancial markets learn about the monthly imports and exports activities of the various industries with a lag and react accordingly.[13]

Furthermore, we test how ¯rms from the same industry that di®er in maturity would react to lagged changes in exchange rates. From their FX exposure, or lack thereof, we may be able to draw conclusions on a ¯rm's ability to hedge over time. The electronic computer

[13] See Guide to Foreign Trade Statistics, Bureau of the Census, U.S. Department of Commerce, section 4, p. 4-1, December 1992.

industry lends itself to this experiment, given that we can construct two industry portfolios which di®er with respect to the age of the ¯rms, while they are comparable in size. In the second panel of table 3, we provide results for the portfolio of the more mature ¯rms for the subperiod 1982-1990. This is to directly compare the exchange rate sensitivities of mature ¯rms to those of the less mature ¯rms for which data only exist for the period 1982-1990. For this period, a lagged change in the exchange rates does not signi¯cantly a®ect the portfolio of mature ¯rms. In the third panel of table 3, we report results for the portfolio of the less mature ¯rms from the same industry. In contrast to those for the mature ¯rms, the equally weighted portfolio (IPLW) and the value weighted portfolio (IPLVW) of the less mature ¯rms, as well as the portfolio of the largest ¯rms (IPLL) exhibit signi¯cant exposures to lagged exchange rates. Again, the portfolio of the smallest ¯rms (IPLS) is not signi¯cantly exposed. In all cases, the signs of the exposures are consistent with our alternative hypothesis (positive), indicating that a lagged appreciation of the dollar signi¯cantly reduced returns.

Our ¯nding that the less mature ¯rms are more signi¯cantly exposed than the mature ¯rms is consistent with our intuition that less mature ¯rms might be engaging in hedging at a smaller scale, or less e®ectively. These results are consistent with the hypothesis that there is learning in hedging and therefore, mature ¯rms manage exposure better than less mature ¯rms. In all cases the contemporaneous change in the exchange rate does not signi¯cantly a®ect equity returns, except for the portfolio of the small mature ¯rms which is signi¯cantly exposed, however, the sign of the exposure (negative) is the opposite from what was hypothesized.

4.3 Robustness Tests

To examine the robustness of our results to the choice of the index, we also use the MG index which is computed by Morgan Guarantee Bank of NY and has been commonly used in earlier studies (e.g., Amihud (1993)). Overall, the exchange rate exposures to the MG movements are very similar to those relative to RX-101. The lagged exchange rate exposure coe±cients of all the alternative industry portfolios have a positive sign indicating that a lagged dollar appreciation reduces industry returns and are statistically signi¯cant (results not reported).

A second issue that may a®ect our estimates of exhange rate exposure is the presence of autocorrelation in the exchange rate index. Generally, autocorrelations are small, ranging from 0.22 (RX-101) (statistically di®erent from zero) to 0.18 (MG). To correct for the presence of autocorrelation in our indices, we use a model where, the unexpected change in the exchange rate index is not the actual change in the index, but the residual in the regression of the change in the index on its lagged change. Again, we obtain similar results: the lagged exchange rate exposures of all the alternative electronic computer industry portfolios are positive, statistically signi¯cant and of similar magnitude with those obtained earlier using the actual changes in the index.

Finally, we also control for possible correlation between the return on the market portfolio and the return on the exchange rate index, using a model similar to (3), but substituting the market factor with an orthogonalized market factor (the orthogonalized market factor is the residual of the regression of the market factor on the exchange rate).[14] A good indication on how correlated the market factor and exchange rates are is given by the adjusted R^2 in the regression of the market factor on the exchange rate factor. Given that the adjusted R^2 are extremely low (0.00063), indicating that the correlation between the market and exchange rates is very low, we should not expect that this issue has a®ected our results in any material way. Indeed, we obtain very similar results to the ones obtained using the original market factor (lagged, positive and statistically signi¯cant exposure) (results not reported).

Overall, the above tests suggest that our results are robust with respect to the choice of exchange rate index, the presence of autocorrelation in the exchange rate index and the use

[14] He, Ng and Wu (1996) use this methodology to test for exchange rate exposure in a sample of Japanese companies and ¯nd that Japanese companies are signi¯cantly a®ected by exchange rate movements.

of the market factor (instead of the orthogonalized market factor).

4.4 FX Exposure at Long-Horizon; Test of H3

First, we test hypothesis **H3** using the Electronic Computer industry returns. We perform long-horizon regressions to capture the long-term exposure of the returns to exchange rate movements. Long-horizon regressions are regressions where the dependent/independent variables are measured over a longer period than the sampling interval. The difference between the return horizon and the sampling interval leads to serial correlation even under the null of no correlation. In our case the sampling interval is one month and the return horizon is 3,6,9 and 12 months. OLS provides consistent estimates, but traditional OLS standard errors cannot be used since the error term is serially correlated. We handle this as in Hansen (1982). The long-horizon estimator of model (4) can be viewed as a generalized method of moment estimator with instruments $x_t = (1, R_{mt}, FX_t)$.

Given that there is still no agreement on what constitutes long-horizon, we provide results using 3, 6, 9 and 12 months. In Figures 1 to 4, we show the one-month exchange rate changes together with the 3, 6, 9 and 12 month exchange rate changes, correspondingly. The long-term trend becomes clear as we move from the one-month to the twelve-month exchange rate changes.

The use of overlapping observations in calculating the long-horizon exposure is necessary in this case, given the fact that floating exchange rates exist since 1973 only, and monthly data for the RX-101 since January 1976. In general, since nonoverlapping data ignores information in the time series, it should produce less efficient estimates than the overlapping data. Given the low autocorrelation (0.22) of the RX-101 exchange rate index, Boudoukh and Richardson (1993) suggest that these efficiency gains are substantial.

The results in Table 4 present the estimates of the long-horizon exposure coefficient γ in model (4) and its standard errors for 3, 6, 9, and 12 month horizon for the entire period, January 1976 to December 1990 and for the three equally divided subperiods. Caution should be exercised when interpreting the subperiod results for 9 and 12 months overlapping observations, since in these tests there are few independent observations. For comparison, we also present results on the one-month (short-horizon) exposure. We ¯nd that the magnitudes of the exposures are generally larger (in absolute value), the longer the horizon. For example, the exposure increases from -0.7779 (one-month) to -1.4254 (12-month) for the subperiod 1981-85 and from -0.1550 to 3.189 for the subperiod 1976-80. The statistical signi¯cance also increases with the horizon. While the Electronic Computer industry is not signi¯cantly a®ected at short horizons, neither for the entire period, nor in any of the subperiods, at long horizons, it is signi¯cantly exposed during the subperiods 1976-80 and 1986-90. More importantly, in both subperiods, the signs of the long-horizon exposures (positive) are consistent with the alternative hypothesis, indicating that an appreciation of the dollar signi¯cantly reduces the industry returns at long horizon.

4.4.1 The long-horizon exposure of the automotive industry

We also examine the long-horizon exposures of the U.S. automotive ¯rms. The automotive industry provides another sample to test whether exchange rate exposures are more pronounced at long horizons. At the same time, we can examine issues related to the e®ect of the dollar/yen movements on the competitiveness of the U.S. automotive ¯rms.

In table 5, we present long-horizon exposure estimates with respect to the RX-101 movements for Ford, GM and Chrysler. For comparison, we also present results on the one-month (short-horizon) exposure. The signs of the FX exposure are largely the same at short and long horizons for the entire period as well as for the di®erent subperiods. The magnitudes of the exposures are generally larger (in absolute value), the longer the horizon. The statistical signi¯cance - with the exception of GM - also increases at long horizons. For example, Chrysler is signi¯cantly exposed in only one subperiod at short horizons (1986-1990), while at long horizons, it is signi¯cantly exposed during 1981-1985 and 1986-1990. Results on long-horizon exposures of Federal and Navistar further support the above conclusions (results are available from the author).[15]

In Table 6, we provide estimates of the long-horizon exposure coe±cients of the automotive ¯rms and the corresponding standard errors for 3, 6, 9 and 12 month-horizon regressions, with respect to the Dollar/Yen exchange rate. These exposures should shed some light on the issue of competitive exposure. At the same time, the use of a single exchange rate alleviates the problem that the ex-ante unknown weights of the exchange rate index may introduce. The three panels present results for Ford, GM and Chrysler. To facilitate the comparison between the long-horizon and the short-horizon (one month) exposures, we also include in the above panels the short-horizon estimates for the Dollar/Yen exposure of Ford, GM and Chrysler. The results from the short and long-horizon regressions di®er signi¯cantly. While in short-horizon regressions, neither Ford, nor GM, nor Chrysler exhibit any signi¯cant exposure to the Dollar/Yen changes in any period, at long-horizons, Ford is signi¯cantly exposed to the Dollar/Yen long-horizon changes for the subperiods 1976-1980 and 1986-1990, GM is signi¯cantly exposed for the subperiod 1981-1985, and Chrysler is signi¯cantly exposed for the subperiods 1981-1985 and 1986-1990. Again, as in the previous tests of long-horizon exposure with respect to the real exchange rate index RX-101, we ¯nd that in general, the magnitudes of the exposures with respect to the Dollar/Yen are larger at long horizons than at short horizons. For example, Ford's one-month exposure increases from 0.206 to 0.589 (12-month exposure) for the entire period (1976-1990), and from 0.381 (0.263) to 0.582 (1.826) for the subperiod 1976-1980 (1986-1990). Very similar results were obtained for the remaining two ¯rms in the automotive industry, Federal and Navistar. While there is no signi¯cant exposure at short horizons, both Federal and Navistar exhibit signi¯cant exposure to the

[15] One additional reason that could explain why exposure is more pronounced at long horizons is that, while short-term hedging was information not publicly available to the market during this period (1976-1990), long-term hedging (e.g., foreign currency borrowing) was. Hence, the ¯nancial markets may be better able to infer the relationship between stock returns and exchange rates at longer horizons. Thanks are due to Michael Adler for suggesting this explanation.

real exchange rates (Dollar/Yen) at long horizons (results are available from the author).

The signs of the long-horizon exposures of the automotive firms to the Dollar/Yen exchange rate are consistent with their overall competitive strategies during this period. As shown in table 6, for the subperiod 1981-1985, Ford's long-horizon exposure is positive whereas both GM's and Chrysler's are negative. Ford's positive exposure indicates that its returns were hurt by the appreciation of the dollar during that period. In contrast, both GM and Chrysler significantly benefited by the dollar's appreciation. GM took advantage of the dollar's large appreciation raising its equity in Isuzu Motors in Japan from 36% to 43% and importing and marketing trucks and midsized cars to the U.S. made by Isuzu and subcompact cars made by Suzuki Motors.[16] Similarly, Chrysler imported 87,500 cars and trucks to the U.S. made by Mitsubishi Motors and marketed them under its own brand name and raised its stake in Mitsubishi in Japan from 15% to 24%.[17] In contrast, Ford's strategic plans were not accomodated by the rising dollar; Ford had started a program to cut the use of foreign components to 5% per car by 1982 from 10% per car in 1971 and despite "threatening to move production overseas, if Japan's imports kept on rising", it never really did so. When Ford decided to jointly manufacture cars with Mazda Motors and KIA in South Korea in 1986, it was a little too late, given that this was going to be a period of a declining dollar.[18]

Finally, the statistical power of the regressions, as measured by the adjusted R^2, is also increasing as the horizon increases. In Table 7, we provide adjusted R^2s for the regressions capturing Chrysler's exchange rate exposure to the Dollar/Yen at different horizons, for the entire period, January 1976 to December 1990 and the three subperiods that we examine. For the entire period, as well as the individual subperiods, the adjusted R^2s increase significantly with the horizon. In particular, for the entire period, the adjusted R^2s increase by 75 percent,

[16]See, e.g., Japan Economy 1/10/84, p10, NY Times 1/18/84, pB9, Forbes 7/30/84, p12.

[17]See, e.g., Asian WSJ 3/12/84, p16, WSJ 4/4/85, p12.

[18]See, e.g., Autonews 4/27/81, p3, Wards Auto 6/82, p10, and WSJ 07/11/86, p37.

from 0.29 for the one-month regression to 0.51 for the 12-month regression. The largest increase in adjusted R^2 occurs in the subperiod 1981-1985 (from 0.22 in the one-month regression to 0.65 in the 12-month regression).

In conclusion, from the investigation of the long-horizon exposures, three points emerge: i) the estimates of the long-horizon exposures have largely the same sign with the estimates for the short-horizon exposures; ii) the magnitudes of the exposures are larger (in absolute value) at long horizons than at short horizons and iii) the statistical signi¯cance of the exposures is much stronger at long horizons than at short horizons.

5 Conclusions

In this paper we investigate the following two reasons as potential explanations for the surprising results of previous research that exchange rate movements do not a®ect ¯rms' values: a) Import activities naturally hedge the exchange rate exposure that arises from the export activities; and b) Exchange rate movements do a®ect ¯rms signi¯cantly at long horizons, even though they may not a®ect them at short horizons.

Our ¯ndings indicate that: i) Exports and imports, the ¯rms' real activities, do a®ect their FX exposure. Firms which belong to the industry with the largest \net exports-to-sales" ratio are systematically mispriced for a long period between 1976 and 1990. In particular, a 1% lagged appreciation of the dollar reduces this industry's returns by 1.09% (signi¯cant at the 5% level). This lagged exposure is consistent with the lagged release of imports and exports data by the U.S. department of Commerce. Controlling for size, less mature ¯rms in this industry are more exposed to a lagged change in exchange rates than mature ¯rms. This may be an indication that hedging is enhanced over time. These results are robust to the use of alternative exchange rate indices and econometric speci¯cations for the market factor and the innovation in the exchange rate index; and ii) In contrast with the lack of signi¯cant FX exposure at short horizons, ¯rms exhibit very signi¯cant exposures, both statistically and economically, at long horizons. The largest net-exporting industry is significantly exposed at long horizons during the subperiods 1976-80 and 1986-90, although it is not affected at short horizons. In addition the sign of its long-horizon exposure (positive) is consistent with our hypothesis that an appreciation of the dollar reduces its returns. Similarly, the U.S. automotive firms are very significantly exposed at long horizons. For example, Chrysler is significantly exposed to the Dollar/Yen movements at long horizons during the subperiods 1981-1985 and 1986-1990, although it is not significantly exposed at short horizons. Moreover, the signs of the long-horizon exposures of the automotive firms are in line with their overall competitive strategies (i.e., imports, exports, foreign acquisitions and cost restructuring). Long-horizon regressions capture the long swings that the dollar has exhibited and reveal the more fundamental long-term exchange rate exposure.

Further research in this area is warranted. The question of which firms should hedge, how (through real hedging (i.e., matching imports and exports) or through financial hedging) and when (i.e., the timing of the hedge), has not been adequately addressed at the theoretical and empirical level. The changes over time in those operations should result in a time-varying exposure. The explicit modeling of the time-variation of the industry FX exposure as a function of the monthly imports and exports is examined in Allayannis (1995). The effect of exchange rate volatility (as opposed to changes in the levels) on equity returns is another issue worth investigating. One could examine the hypothesis that if firms' success in hedging FX risk is inversely related to exchange rate volatility, (since it is harder to hedge effectively when exchange rates are more volatile), then one should expect to find a negative relationship between exposure and exchange rate volatility.

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Table 1

OLS estimates for the largest net exporter: Computer Electronics Industry; Test of H1

The table provides parameter estimates for the model specified by equation (1)

$$R_{it} = \beta_0 + \beta_1 R_{mt} + \beta_2 FXI_t + \epsilon_t; \; t = 1; \ldots; T \tag{6}$$

for the Industry at the 4-SIC digit with the largest Net Export to Sales ratio (SIC 3571 : Computer Electronics). The estimates (top) and the corresponding t-statistics (bottom) for the constant (β_0), the real return on the CRSP value weighted portfolio (β_1) and the changes in the real exchange rate index (β_2) are presented. Adjusted R^2 are also shown. The period covered is January 1976 to December 1990 and subperiod results are also provided. The data frequency is monthly.

Period	β_0	β_1	β_2	R^2
1976-1990	0.0094	1.8149	-0.4573	0.42
	(1.3018)	(11.5020)	(0.7670)	
1976-1980	0.0467	2.5825	-0.1550	0.57
	(3.6961)	(8.8085)	(0.1084)	
1981-1985	-0.0129	1.6030	-0.7779	0.36
	(1.1379)	(5.9889)	(0.9709)	
1986-1990	-0.0051	1.4445	-0.5935	0.40
	(0.4113)	(6.4040)	(0.6151)	

Table 2

SUR estimates for the individual firms in the Computer Electronics Industry

This table presents SUR estimates for the individual firms in the Computer Electronic Industry according to model (2).

$$R_{it} = \beta_{0i} + \beta_{1i} R_{mt} + \beta_{2i} FXI_t + \epsilon_{it} \quad i = 1, \ldots, 5; \quad t = 1, \ldots, T \tag{7}$$

The estimates of the changes of exchange rates for each individual firm β_{2i} (top) and its t-statistic (bottom) are shown. In the second panel, we also report the $\chi^2(5)$ statistic that examines the joint test that all five slope coefficients (β_{2i}) are zero. The period covered is January 1976 to December 1990 and results for 5 subperiods are also given. Data frequency is monthly.

$Firm_i$	1976 – 1990	1976 – 1978	1979 – 1981	1982 – 1984	1985 – 1987	1988 – 1990
Tandy	0.185	0.764	-0.121	1.975	-2.304	0.691
	(0.320)	(0.386)	(0.079)	(1.282)	$(2.615)^{*}$	(0.818)
Commodore	1.606	-1.800	-0.507	2.126	2.778	3.457
	$(1.507)^{**}$	(0.474)	(0.203)	(1.005)	$(1.607)^{**}$	(1.425)
Alpha	-0.9981	-6.938	-0.021	-0.026	-2.499	-0.618
	(1.326)	$(2.600)^{*}$	(0.014)	(0.017)	$(1.827)^{*}$	(0.412)
Qantel	-1.927	-3.278	-1.739	-1.405	-1.575	-4.463
	$(1.620)^{**}$	(1.024)	(0.899)	(0.731)	(0.716)	(1.028)
Electronic	-0.195	1.721	1.130	-0.356	-1.310	-3.403
	(0.205)	(0.416)	(0.637)	(0.148)	(0.866)	$(2.574)^{*}$

* significant for one-tailed test at 5 percent level

** significant for one-tailed test at 10 percent level

	1976 ¡ 1990	1976 ¡ 1978	1979 ¡ 1981	1982 ¡ 1984	1985 ¡ 1987	1988 ¡ 1990
Â^{2}(5)	8.388	8.3188	1.7214	2.7612	8.9500	14.578
sign.level	0.1360	0.1395	0.8861	0.7360	0.1109	0.0123

Table 3

OLS estimates for Computer Electronic Industry including lagged variable; Test of H2 (using the RX-101 index);

This table provides parameters estimates for model(3)

$$R_{it} = \alpha_i + \beta_i R_{mt} + \delta_{1i} FXI_t + \delta_{2i} FXI_{t-1} + \epsilon_t; \quad t = 1; \ldots; T \qquad (8)$$

investigating possible mispricing using the **RX-101** index; The estimates for the constant (α_i), the market, (β_i), the change in the exchange rate (δ_{i1}) and the lagged change in the exchange rate (δ_{i2}) are shown in the top panel for the equally weighted (value weighted) industry portfolio of the more mature firms IPMW (IPMVW), (those for which data exist between 1976-1990) and two size portfolios, (large (IPML) and small (IPMS)). In the second panel, estimates for the same parameters are provided for the period 1982-1990. In the third panel, coefficient estimates are provided for a different set of firms in the industry, namely those for which data exist between 1982-1990 (Less Mature). In each case, t-statistics are reported in parentheses underneath the coefficient estimates. Size portfolio results are also shown. The data frequency is monthly.

1976-1990

Portfolio	α_i	β_i	δ_{i1}	δ_{i2}	R^2
IPMW	0.009	1.664	-0.348	1.097	0.45
	(1.397)	(11.970)**	(0.655)	(2.066)*	
IPMVW	0.002	0.324	0.032	0.177	0.42
	(1.795)	(11.286)**	(0.300)	(1.615)	
IPML	0.023	1.693	0.590	1.270	0.35
	(2.642)**	(9.663)**	(0.881)	(1.897)*	
IPMS	0.003	1.635	-1.286	0.924	0.35
	(0.451)	(9.890)**	(2.037)*	(1.464)	

*signi¯cant for one-tailed test at 5 percent level

**signi¯cant at 5 percent level

1982-1990

Portfolio	®$_i$	¯$_i$	±$_{i1}$	±$_{i2}$	R^2
IPMW	-0.012	1.469	-0.545	0.735	0.46
	(1.446)	(9:657)¤¤	(0.955)	(1.287)	
IPMVW	-0.001	0.297	0.011	0.107	0.46
	(0.541)	(9:186)¤¤	(0.097)	(0.882)	
IPML	0.004	1.482	0.657	1.098	0.34
	(0.401)	(7:368)¤¤	(0.871)	(1.454)	
IPMS	-0.028	1.456	-1.748	0.372	0.36
	(2.772)	(7:751)¤¤	(2:481)¤	(0.528)	

*signi¯cant for one-tailed test at 5 percent level

**signi¯cant at 5 percent level

1982-1990

Portfolio	®$_i$	¯$_i$	±$_{i1}$	±$_{i2}$	R^2
IPLW	-0.014	1.592	-0.731	1.729	0.43
	(1.387)	(8:856)¤¤	(1.084)	(2:567)¤	
IPLVW	0.002	0.409	-0.001	0.522	0.48
	(1.052)	(9:657)¤¤	(0.008)	(3:288)¤	
IPLL	0.009	1.656	-0.099	1.929	0.50
	(1.007)	(10:004)¤¤	(0.160)	(3:111)¤	
IPLS	0.036	1.528	-1.361	1.529	0.19
	(2.257)	(5:180)¤¤	(1.231)	(1.383)	

* signi¯cant for one-tailed test at 5 percent

** signi¯cant at 5 percent

Table 4

Long-Horizon Regressions; Test of H3

This table provides estimates for the coefficient γ_J with standard errors in parenthesis, to the change in the RX-101 exchange rate index, for the model specified by equation (4).

$$\sum_{i=1}^{J} R_{t+i} = \alpha_J + \beta_J \sum_{i=1}^{J} R_{mt} + \gamma_J \sum_{i=1}^{J} FXI_t + \epsilon_t(J); \quad t = 1, \ldots, T \tag{9}$$

We use the real returns for the largest net exporting industry at the 4-digit SIC level (Electronic Computers), the CRSP value weighted index adjusted for inflation and the RX-101 exchange rate index. The table presents results for horizons of 1 (short horizon), 3, 6, 9 and 12 months for the entire period (January 1976 to December 1990) and for three subperiods.

Electronic Computers Industry

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	-0.4573	-0.1550	-0.7779	-0.5935
	(0.5962)	(1.4298)	(0.8012)	(0.9648)
3	-0.1550	1.0182	-0.4490	-0.6473
	(0.5147)	(1.8783)	(0.8231)	(0.5229)
6	0.1626	3.2072	-0.6129	0.43185
	(0.8608)	(1.9920)	(1.0525)	(0.6297)
9	0.1822	3.4224^{**}	-1.0286	2.1578^{*}
	(1.2490)	(1.9026)	(1.1041)	(0.5805)
12	-0.1478	3.189^{**}	-1.4254	1.9284^{*}
	(1.4540)	(1.7597)	(1.3303)	(0.6197)

*significant at the 5 percent

** significant at the 10 percent

Table 5

Long-Horizon Regressions (RX101); Test of H3

This table provides estimates for the coefficient γ_J with standard errors in parenthesis, of the change in exchange rate index, for the model specified by equation (4).

$$\sum_{i=1}^{J} R_{t+i} = \alpha_J + \beta_J \sum_{i=1}^{J} R_{mt+i} + \gamma_J \sum_{i=1}^{J} FXI_{t+i} + \epsilon_t(J); \; t = 1; \ldots; T \qquad (10)$$

We use the real returns for FORD (GM and Chrysler), the CRSP value weighted index adjusted for inflation and the real exchange rate index RX-101. The table presents results for horizons of 1 (short horizon), 3, 6, 9 and 12 months for the entire period (January 1976 to December 1990) and for three subperiods.

Using RX-101

FORD

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.361	1.663**	-0.383	1.161*
	(0.4163)	(0.9213)	(0.7495)	(0.5652)
3	0.1353	2.5145	-0.8084	1.5293**
	(0.7429)	(1.7908)	(0.8670)	(0.7793)
6	0.0176	2.4765*	-0.5463	2.4441*
	(0.9221)	(1.1529)	(0.8133)	(0.9210)
9	0.5075	1.7191*	-0.0837	4.5492*
	(1.1330)	(0.6432)	(0.7506)	(0.5568)
12	0.8568	2.4849*	0.0559	5.2197*
	(1.1873)	(0.6340)	(0.7312)	(0.4543)

*significant at the 5 percent

** significant at the 10 percent

GM

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.0582	1:156¤¤	¡ 0:9415¤¤	0:9541¤¤
	(0.3095)	(0.6020)	(0.5268)	(0.5127)
3	-0.2595	1.7360	¡ 1:3416¤	0.8723
	(0.4319)	(1.0413)	(0.4659)	(0.5529)
6	-0.6930	0.0181	¡ 1:6256¤	0.5103
	(0.4983)	(0.7954)	(0.4758)	(0.6824)
9	-0.6517	-0.2470	¡ 1:6190¤	0.1032
	(0.4325)	(0.4586)	(0.4591)	(0.3422)
12	-0.4654	0.5694	¡ 1:5790¤	0.1140
	(0.5203)	(0.3713)	(0.4261)	(0.4906)

CHRYSLER

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.3248	3:0074¤¤	0.1140	0.4367
	(0.6919)	(1.7214)	(1.2284)	(0.8369)
3	-0.7266	3.2430	-1.7146	1:3516¤¤
	(0.9809)	(2.7411)	(1.5517)	(0.7324)
6	-1.3721	1.7544	¡ 2:3398¤¤	1:9304¤¤
	(1.1942)	(1.5631)	(1.2337)	(1.0721)
9	-1.5399	0.5098	¡ 2:7108¤	3:7366¤
	(1.4627)	(1.0275)	(1.3857)	(1.1437)
12	-1.7066	1.0655	¡ 3:2370¤	4:2279¤
	(1.6852)	(0.7300)	(1.5311)	(0.9919)

*signi¯cant at the 5 percent

** signi¯cant at the 10 percent

Table 6

Long-Horizon Regressions (USD/JYEN); Test of H3

This table provides estimates for the coefficient γ_J with standard errors in parenthesis, of the change in the USDollar/JYen exchange rate, for the model specified by equation (4).

$$\sum_{i=1}^{J} R_{t+i} = \alpha_J + \beta_J \sum_{i=1}^{J} R_{mt} + \gamma_J \sum_{i=1}^{J} FXI_t + \epsilon_t(J); \quad t = 1, \ldots, T \tag{11}$$

We use the real returns for FORD (GM and Chrysler), the CRSP value weighted index adjusted for inflation and the USDollar/JYen exchange rate. The table presents results for horizons of 1 (short horizon), 3, 6, 9 and 12 months for the entire period (January 1976 to December 1990) and for three subperiods.

Using Dollar/Yen

FORD

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.2060	0.3810	0.1010	0.2630
	(0.1598)	(0.2719)	(0.3300)	(0.2139)
3	-0.0036	0.1804	-0.1654	0.3304
	(0.2341)	(0.3538)	(0.3147)	(0.3625)
6	0.1531	0.3609	-0.0234	0.5625
	(0.3193)	(0.3693)	(0.3449)	(0.4451)
9	0.4091	0.4719^{*}	0.0888	1.2868^{*}
	(0.3457)	(0.2488)	(0.3149)	(0.4111)
12	0.5893^{**}	0.5826^{*}	0.0488	1.8260^{*}
	(0.3067)	(0.1822)	(0.3211)	(0.1973)

*significant at the 5 percent

** significant at the 10 percent

GM

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.0650	0.1940	-0.2300	0.2750
	(0.1224)	(0.1797)	(0.2361)	(0.1925)
3	-0.1458	0.1187	-0.5123*	0.1013
	(0.1317)	(0.1956)	(0.2016)	(0.2411)
6	-0.2036	0.0886	-0.5332*	-0.1443
	(0.1620)	(0.1985)	(0.2134)	(0.3113)
9	-0.1715	0.0799	-0.5543*	-0.2065
	(0.1513)	(0.1305)	(0.2022)	(0.1655)
12	-0.1255	0.1624	-0.6026*	-0.0591
	(0.1594)	(0.0970)	(0.1839)	(0.2402)

CHRYSLER

Horizon J	1976-1990	1976-1980	1981-1985	1986-1990
1	0.1000	0.0400	0.2540	0.1980
	(0.2659)	(0.5128)	(0.5415)	(0.3088)
3	-0.1307	-0.0350	-0.4880	0.5964*
	(0.2899)	(0.3044)	(0.5717)	(0.2492)
6	-0.0651	0.0436	-0.7678**	0.8299*
	(0.4061)	(0.2477)	(0.4683)	(0.3437)
9	-0.0641	0.0353	-1.0567*	1.4456*
	(0.5074)	(0.2587)	(0.5483)	(0.3868)
12	-0.0857	0.0962	-1.3374*	1.6958*
	(0.5758)	(0.2150)	(0.6254)	(0.3035)

*significant at the 5 percent

** signi¯cant at the 10 percent

Table 7

Long-Horizon Regressions (USD/JYEN)

Adjusted R^2 for di®erent horizons using Dollar/Yen

This table provides adjusted R^2 for the model speci¯ed by equation (4).

$$\sum_{i=1} R_{t+i} = \alpha_J + \beta_J \sum_{i=1} R_{mt} + \gamma_J \sum_{i=1} FXI_t + \epsilon_t(J); \quad t = 1; ...T \tag{12}$$

We use the real returns for Chrysler, the CRSP value weighted index adjusted for in°ation and the USDollar/JYen exchange rate. The table presents R^2s for horizons of 1, 3, 6, 9 and 12 months for the entire period (January 1976 to December 1990) and for three subperiods.

CHRYSLER	1MON	3MON	6MON	9MON	12MON
1976-1990	0.290	0.375	0.423	0.472	0.514
1976-1980	0.240	0.220	0.167	0.166	0.429
1981-1985	0.220	0.296	0.512	0.610	0.654
1986-1990	0.570	0.745	0.697	0.781	0.831