Accounting for U.S. Real Exchange Rate Changes

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This study measures the proportion of U.S. real exchange rate movements that can be accounted for by movements in the relative prices of nontraded goods. The decomposition is done at all possible horizons that the data allow—from one month up to 30 years. The accounting is performed with five different measures of nontraded-goods prices and real exchange rates, for exchange rates of the United States relative to a number of other high-income countries in each case. The outcome is surprising: relative prices of nontraded goods appear to account for almost none of the movement of U.S. real exchange rates. Special attention is paid to the U.S. real exchange rate with Japan. The possibility of mismeasurement of traded-goods prices is explored.

The real exchange rate is a measure of one country's overall price level relative to another country's. It is often associated with the price of nontraded goods relative to traded goods. To see why, consider a price index for a country that is a geometric weighted average of traded- and nontraded-goods prices:

$$p_t = (1 - \alpha)p_t^T + \alpha p_t^N,$$

I thank participants at seminars at the Federal Reserve Bank of Kansas City, the NBER, the University of Washington, and the Castor workshop, as well as the referee at this *Journal* for useful comments. Mike Hendrickson and Avery Tillett-Ke performed outstanding research assistance. Part of the work for this project was completed while I was a visiting scholar at the Federal Reserve Bank of Kansas City. The views expressed in this paper do not represent those of the Federal Reserve Bank of Kansas City or the Federal Reserve System. The National Science Foundation provided support for this project under a grant to the NBER.

[*Journal of Political Economy*, 1999, vol. 107, no. 3]
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where p_t is the log of the price index, p_t^T is the log of the traded-goods price index, p_t^N is the log of the nontraded-goods price index, and α is the share that nontraded goods take in the price index. Letting an asterisk represent the foreign country, one can also write

$$p_t^* = (1 - \beta)p_t^{T*} + \beta p_t^{N*},$$

where β is nontraded goods' share in the foreign price index. Then the real exchange rate is given by

$$q_t = x_t + y_t, \tag{1}$$

where

$$q_t = s_t + p_t^* - p_t,$$
$$x_t = s_t + p_t^{T*} - p_t^T,$$
$$y_t = \beta(p_t^{N*} - p_t^{T*}) - \alpha(p_t^N - p_t^T).$$

Here, s_t is the log of the domestic currency price of foreign currency.

Equation (1) indicates that the log of the real exchange rate is composed of two parts: the relative price of traded goods between the countries, x_t, and a component that is a weighted difference of the relative price of nontraded- to traded-goods prices in each country, y_t.

The real exchange rate for the United States relative to other high-income countries has fluctuated dramatically over the past 25 years. The short-term and longer-term movements have been much greater than what the United States witnessed over the 25-year period immediately preceding that. Most of the recent theoretical literature on real exchange rates has emphasized movements in the nontraded-goods component, y_t.[1] This study is an attempt to measure the significance of the nontraded-goods component in U.S. real exchange rate movements.

Determining precise price indexes for nontraded goods and traded goods is an impossible task given the quality of data available. Still, since movements in the relative price of nontraded goods play such a large role in real exchange rate theory, it seems a worthwhile task to attempt some measure of this price. In this paper, I look at five measures of nontraded-goods prices. The first is based on consumer price indexes (CPIs); it classifies services and housing as

[1] Recent examples include Asea and Mendoza (1994), Brock (1994), Brock and Turnovsky (1994), De Gregorio, Giovannini, and Krueger (1994), Samuelson (1994), De Gregorio and Wolf (1994), Razin (1995), and Obstfeld and Rogoff (1996).

nontradable and commodities as tradable. The second is constructed from an OECD database of output prices that previous studies have used to produce traded- and nontraded-goods price indexes (see De Gregorio, Giovannini, and Wolf 1994; De Gregorio and Wolf 1994; Canzoneri, Cumby, and Diba 1996). The third uses price deflators for personal consumption expenditures, letting the price index for expenditure on goods be the measure of traded-goods prices and the price index for services be the measure of nontraded-goods prices.[2] The fourth measure is a cruder one: the aggregate producer price index (PPI) is used as the measure of traded-goods prices, and the nontraded component, y_t, is constructed from aggregate CPIs relative to aggregate PPIs. Finally, it has been argued that a large component of consumer prices consists of nontraded marketing services.[3] The behavior of prices for 36 marketing and distribution services relative to the general price level in Japan is investigated.

This is not the first study to construct traded and nontraded price indexes (see, e.g., De Gregorio, Giovannini, and Wolf 1994; De Gregorio and Wolf 1994; Kakkar and Ogaki 1994; Stockman and Tesar 1995; Canzoneri et al. 1996). However, there are two serious problems with earlier work. Many of the previous studies look only at the price of nontraded to traded goods for an individual country. But the nontraded-goods component, y_t, is a relative relative price: the relative price of nontradables to tradables in one country relative to that relative price in another country. So simply looking at $p_t^N - p_t^T$ will not reveal much about q_t. We must look at how $\alpha(p_t^N - p_t^T)$ moves relative to $\beta(p_t^{N*} - p_t^{T*})$. For example, Marston (1987) finds that the relative prices of nontraded goods in each of Japan and the United States are related to the relative average labor productivity in each country. De Gregorio, Giovannini, and Krueger (1994) and De Gregorio, Giovannini, and Wolf (1994) look at determinants of $p_t^N - p_t^T$ such as productivity, fiscal expenditures, and private expenditures for a large number of countries. Asea and Mendoza (1994) relate the relative price of nontradables to the relative capital intensities of traded and nontraded sectors, using data for 14 countries. In none of these studies do the authors examine the behavior of $p_t^N - p_t^T$ relative to $p_t^{N*} - p_t^{T*}$.

The second problem is that these studies do not examine how important the nontraded price component is in the overall move-

[2] This measure has been used by Kakkar and Ogaki (1994) and Stockman and Tesar (1995).

[3] This argument is expounded in Engel (1993), Froot and Rogoff (1995), and Engel and Rogers (1996).

ment of the real exchange rate. That is, they do not compare the contribution of movements in y_t to movements in x_t in determining the overall changes in the real exchange rate.[4]

Some theories of real exchange rate movements distinguish short-run from long-run behavior. For example, if nominal prices are sticky and nominal exchange rates fluctuate greatly, then short-run movements of the real exchange rate can be dominated by movements in the tradable component, x_t. As prices adjust over time, the relative price of the nontraded-goods component, y_t, becomes more important.[5] In examining the relative size of the x_t and y_t components, one must therefore look at different horizons. This study examines the average change and the variance of the change in these components at all horizons that the data allow, in some cases from horizons as short as one month to as long as 30 years.

The impression from the evidence presented here is that the nontraded-goods component has accounted for little of the movement in the real exchange rates for the United States at any horizon. While I cannot be very confident about my findings at longer horizons, it appears that for short and medium horizons, knowledge of the behavior of the relative price on nontraded goods contributes practically nothing to one's understanding of U.S. real exchange rates.

I. Consumer Prices

This section measures real exchange rates using consumer prices using monthly OECD data from January 1962 to December 1995 for Canada, France, Germany, Italy, Japan, and the United States. A traded-goods price index for each country is constructed from the food and all goods less food subindexes, and a nontraded-goods price index is built from the shelter and all services less shelter subindexes.[6]

These are imperfect measures of traded- and nontraded-goods prices. First, it is dubious that some of the items in the traded-goods category are actually traded, for example, restaurant meals. In general, consumer prices probably should be thought of as the price of a joint product: the good itself and the service that brings the good to market. While the good might be tradable, the marketing service is probably closer to being nontradable. Conversely, some of the

[4] Rogers and Jenkins (1995) do decompose the variance of the real exchange rate at four horizons into a component attributable to food prices and a component attributable to all other goods.

[5] See Froot and Rogoff (1995) and Rogoff (1996) for recent surveys of this literature.

[6] See App. A for more details on the construction of these series.

items contained in the nontraded category are probably tradable. Many services, such as financial services, are traded. The shelter category includes not only rent on dwellings, which is probably legitimately nontraded, but also, for example, the cost of heating the dwelling, which has a large tradable component. These measurement error issues are explicitly addressed in Section VI.

The drift and the variance of a variable measure two different notions of movement in that variable over time. The mean-squared error (MSE) of the change in the real exchange rate—which is the sum of the squared drift and the variance—is a comprehensive measure of movement. I shall measure how much of the MSE of changes in q_t is attributable to changes in x_t.

The decomposition of the MSE of q_t into its x_t and y_t components depends on the comovements of x_t and y_t. In practice, for all the measures of x_t and y_t except the ones described in Section IV, the two series are nearly uncorrelated in first differences. Appendix B discusses these issues in detail and explains how the drift, variance, and MSEs are calculated.

Why do I examine changes in q_t (and x_t and y_t) rather than the levels of the series? I cannot reject the null of a unit root at the 5 percent level in any of these three series using an augmented Dickey-Fuller test.[7] All three of these series are highly persistent. The series are probably too short to distinguish between highly persistent but stationary series and series that have a unit root. If shocks to tastes and technology are permanent, then y_t may not be stationary. If deviations from purchasing power parity (PPP) for traded goods are transitory, then x_t will be stationary.[8] Thus one expects the variance of k-differences of x_t to converge as k gets large but the variance of k-differences of y_t to grow linearly with k. But we shall see that in the relatively short time series, the behavior of x_t and y_t is indistinguishable from the behavior of two independent random walks.

The decompositions are presented graphically. Figure 1 shows the MSE decomposition using all 408 months of the data set. One can calculate the MSEs of k-differences up to a horizon of $k = 406$. The horizon appears on the horizontal axis of the graphs. Plotted is the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component, $x_{t+k} - x_t$.

If deviations from PPP in traded goods are transitory, the share of the MSE accounted for by the x_t component would be larger at

[7] Indeed, I fail to reject a unit root for every measure of q_t, x_t, and y_t for every country in this study.

[8] Engel (1996) emphasizes that tests that have rejected unit roots in q_t might fail to detect a y_t component that has a unit root but a relatively small innovation variance (compared to that of x_t).

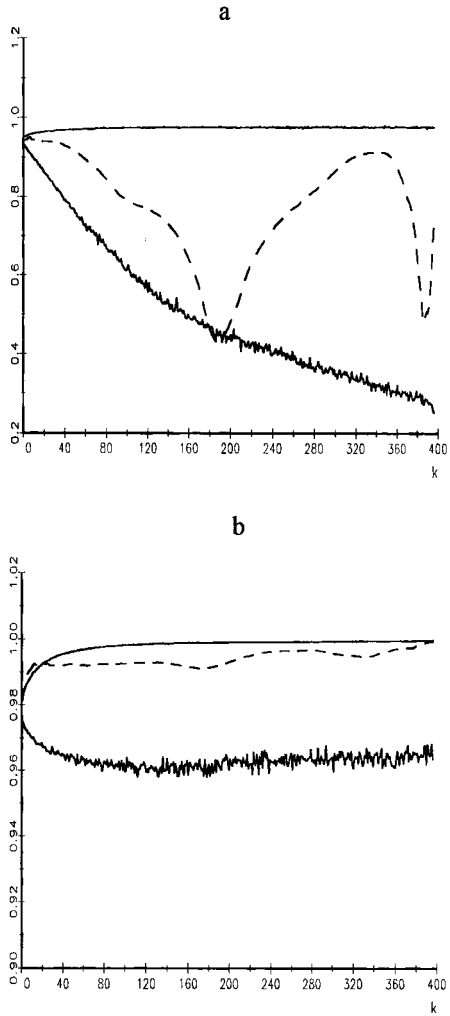

FIG. 1.—MSE decomposition: CPI data. MSE ratio estimates and 95 percent confidence intervals: $\mathrm{MSE}(x_{t+k} - x_t)/[\mathrm{MSE}(x_{t+k} - x_t) + \mathrm{MSE}(y_{t+k} - y_t)]$, January 1962 to December 1995. *a*, Canadian/U.S. real exchange rate. *b*, Japanese/U.S. real exchange rate. *c*, France, Germany, and Italy.

short horizons. So the lines plotted in the graphs should be declining as the horizon increases. However, we should note that, at long horizons, very few observations are used in the calculation of the MSE decomposition, so the statistic is not likely to be very reliable.

Figure 1 is striking in that it shows that the x_t component (the traded-goods component) accounts for nearly 100 percent of the

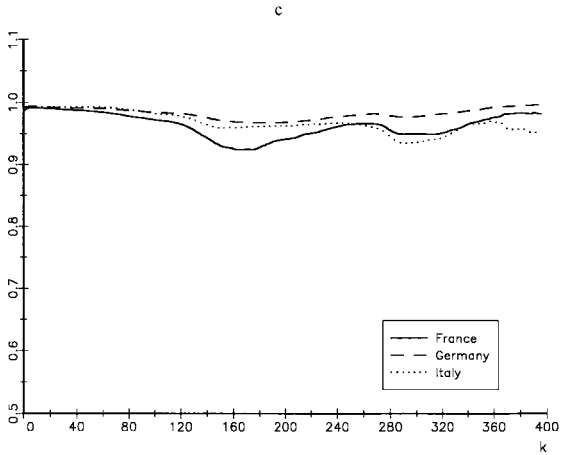

FIG. 1.—(*Continued*)

MSE of the U.S. real exchange rate changes at all horizons, for all rates except the Canadian-U.S. rate. It is notable that there is no apparent decline in the share of the MSE accounted for by x_t even as the horizon increases.

At short horizons for the Canadian-U.S. rate, the traded-goods component accounts for over 95 percent of the movement; but as the horizon increases, the importance of the x_t component drops—to about 45 percent at the 15-year horizon—but the component begins to be more important again at horizons past 15 years. The longer-horizon numbers, however, are less reliable (although 168 monthly observations are available here to calculate the 20-year MSEs).

Figures 1*a* and 1*b* present 95 percent confidence intervals for the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component, $x_{t+k} - x_t$ for the U.S.-Canadian (1*a*) and U.S.-Japanese (1*b*) real exchange rates. The confidence intervals are calculated from Monte Carlo experiments (as described in App. B). The confidence intervals are calculated under the null that $x_{t+1} - x_t$ and $y_{t+1} - y_t$ are independent, normally distributed random variables with means and variances equal to their sample means and variances.

Note that the actual ratio is within the 95 percent confidence band at all horizons. I can give a statistical interpretation of this result: one cannot reject the null that x_t and y_t are independent random walks. Thus there is nothing in the data that suggests that the importance of the x_t component in explaining movements in q_t diminishes over time (or that the importance of the y_t component increases). This should not be taken as conclusive evidence that the x_t component is not stationary. With only 34 years of data, as I have already

noted, one cannot develop conclusive evidence of the stationarity or nonstationarity of x_t.

I do not present the confidence intervals for the real exchange rates of the United States relative to France, Germany, or Italy in order to save space and have graphs that are legible. However, the same results found in the Canadian and Japanese cases carry over to these real exchange rates: the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component lies within the 95 percent confidence interval at all horizons.

It is interesting to compare these findings with those of Kakkar and Ogaki (1994). These authors construct a y_t variable and test for cointegration of that variable with q_t. Their measures of y_t and q_t correspond to the ones used in this paper in Sections III and IV. They find that they cannot reject the null that y_t and q_t are cointegrated using Park's (1990) test for the null of stochastic cointegration. This finding holds for a number of different real U.S. exchange rates over a variety of time periods. If y_t and q_t are cointegrated, then x_t and y_t cannot be independent random walks. But Kakkar and Ogaki's results are not necessarily inconsistent with the ones shown here. I fail to reject the null that x_t and y_t are independent random walks, whereas they fail to reject the null that y_t and q_t are cointegrated: both failures to reject likely reflect the low power of tests to distinguish between unit roots and stationarity in relatively short time spans.

The finding that over 95 percent of the MSE of the real exchange rate is accounted for by the MSE of the traded-goods component holds if we examine the variance rather than the MSE for each of the real exchange rates. It is also the case that the drift in q_t is almost entirely due to the drift in x_t for all the real exchange rates except the U.S.-Canadian rate. The portion of the drift of q_t attributable to the drift of x_t is .485 for the U.S.-Canadian rate, .993 for the U.S.-French rate, .996 for the U.S.-German rate, .857 for the U.S.-Italian rate, and .999 for the U.S.-Japanese rate.

As Mussa (1986) and others have noted, the variance of the U.S. real exchange rate has been much lower in fixed nominal exchange rate periods than in floating exchange rate periods. Figure 1 uses data from both fixed and floating periods for the United States. One might suspect that the results would look different from those from just the fixed-rate period, but that is not true. Figure 2*a* shows the MSE decomposition for the floating rate period, 1973–95, and figure 2*b* shows the same decomposition for the fixed-rate period, 1962–69. In both cases, at all horizons, nearly all of the movement in the real exchange rate comes from movements in the traded-goods component.

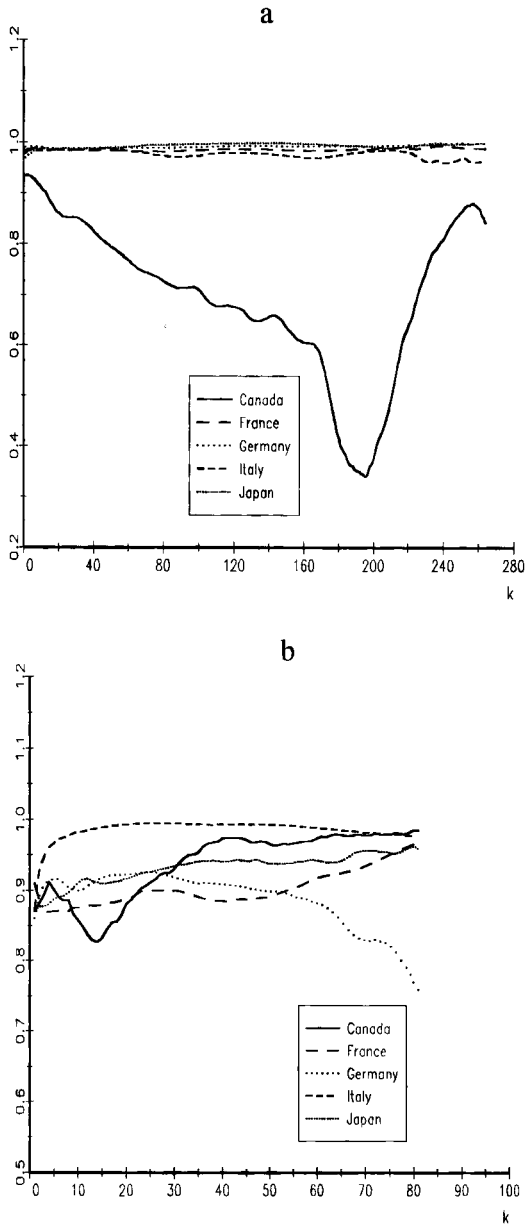

FIG. 2.—MSE decomposition: CPI data. $\mathrm{MSE}(x_{t+k} - x_t)/[\mathrm{MSE}(x_{t+k} - x_t) + \mathrm{MSE}(y_{t+k} - y_t)]$. *a*, January 1973 to December 1995. *b*, January 1962 to September 1969.

The fact that the traded-goods component accounts for nearly all of the movement in real exchange rates even when the nominal exchange rate is fixed is reminiscent of the finding of Froot, Kim, and Rogoff (1995). Using 700 years of data, they find that the volatility of deviations from the law of one price is no greater in the twentieth century under floating exchange rates than it was in previous centuries when the exchange rate was essentially fixed.

II. Output Prices

This section uses output prices from the OECD's international sectoral database. Annual price indexes for 19 categories of output are calculated by dividing nominal output figures by their real counterparts.[9] De Gregorio, Giovannini, and Wolf (1994) classify these sectors as tradable or nontradable using data on the export share in total production. They conclude that all the commodities are tradable. All services are categorized as nontradable, except for the transport, storage, and communication category. The traded and nontraded price indexes used here are constructed as in De Gregorio, Giovannini, and Wolf.

Although data are available for 14 countries, because of missing observations, it was possible to construct long time series of real exchange rates for only seven countries relative to the United States: Canada, Germany, Denmark, Finland, France, Japan, and Norway. The starting dates for these data vary from country to country. In some cases the data start in 1960 (Finland, Germany, and the United States); in others the data begin as late as 1970 (France and Japan). The Norwegian data end in 1988, the U.S. and Canadian data end in 1989, and for the other countries the data end in 1990.

Figure 3 shows the MSE decomposition using all the data. These graphs tell almost exactly the same story as figure 1 tells using consumer price data. At all horizons, for all countries with the exception of Canada, over 90 percent of the MSE of real exchange rate changes is accounted for by the MSE of the traded-goods component, x_t. There is no tendency for this component to become less important as the horizon lengthens.

As with the CPI data, for Canada the traded-goods share accounts for a large share of real exchange rate movements at short horizons:

[9] The sectors are agriculture; mining; food, beverages, and tobacco; textiles; wood and wood products; paper, printing, and publishing; chemicals; nonmetallic mineral products; basic metal products; machinery and equipment; other manufactured products; electricity, gas, and water; construction; wholesale and retail trade; restaurants and hotels; transport, storage, and communications; finance, insurance, and real estate; community, social, and personal services; and government services.

over 80 percent of 1-year changes. This drops to around 50 percent at the 16-year horizon but then begins to rise again. Of course, very few observations are used to calculate the longer-horizon numbers.

Figures 3*a* and 3*b* show the confidence intervals for the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component, $x_{t+k} - x_t$, for the U.S.-Canadian and U.S.-Japanese real exchange rates, respectively. These confidence intervals are constructed in the same way as those of the previous section. It is notable that for Canada, the confidence intervals are very wide at the longer horizons. As was the case in the previous section, the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component lies within the confidence bands at all horizons. This can be interpreted as a failure to reject the hypothesis that x_t and y_t are independent random walks. I have not shown the confidence intervals for the other real exchange rates because of space considerations, but the same pattern holds true for them: the actual ratio of the MSE of the traded-goods component to the total MSE of the real exchange rate lies within the confidence bands at all horizons.

As was true with the CPI data, the variance decompositions look very much like the MSE decompositions for all exchange rates. For all countries except Canada, the traded-goods component also accounts for over 90 percent of the drift.

Figure 4 shows the MSE decomposition for the years 1973–89: the floating nominal exchange rate period. For all countries except Canada, the traded-goods component is responsible for over 95 percent of the MSE of the real exchange rate changes at all but the longest horizons. For Canada, this fraction is around 85 percent for all horizons.

The CPI and output price data reach the same conclusion: that very little real exchange rate movements for the United States are attributable to the relative price of nontraded goods. This suggests that the results for the CPI data do not arise from the fact that consumer prices measure the price of both the tradable product and the nontradable service that brings the good to market. The marketing service is much less important in the output price data, yet it shows the same pattern as the CPI data.

III. Personal Consumption Deflators

This section uses data from national income accounts for personal consumption deflators. The price indexes are constructed by dividing nominal consumption expenditures by real consumption expenditures. The traded-goods price index is the deflator for personal consumption expenditure on commodities, whereas the nontraded

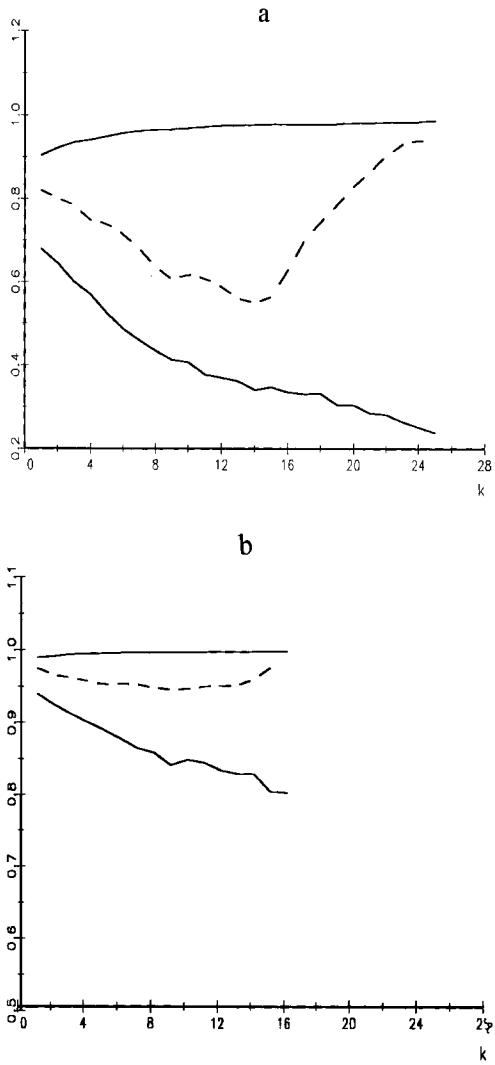

FIG. 3.—MSE decomposition: output prices. MSE ratio estimates and 95 percent confidence intervals: $\mathrm{MSE}(x_{t+k} - x_t)/[\mathrm{MSE}(x_{t+k} - x_t) + \mathrm{MSE}(y_{t+k} - y_t)]$. *a*, Canadian/U.S. output price real exchange rate, 1960–89. *b*, Japanese/U.S. output price real exchange rate, 1970–89. *c*, Germany, Denmark, Finland, France, and Norway, 1960–89.

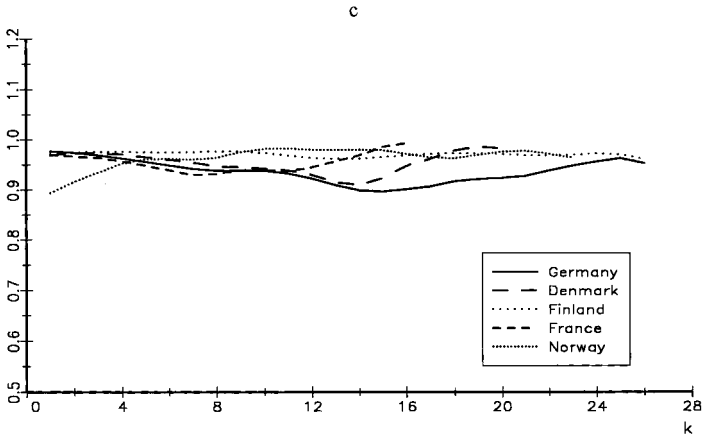

FIG. 3.—(*Continued*)

index is the deflator for personal consumption expenditure on services. This is the categorization used by Stockman and Tesar (1995) for these data. Data are quarterly for Canada, France, Italy, Japan, Sweden, the United Kingdom, and the United States. The data start in the first quarter of 1970 (except those for Sweden, which start in the first quarter of 1980).

Figure 5 displays the MSE decomposition. Again, for all real exchange rates (with the exception of the U.S.-U.K. rate) at all hori-

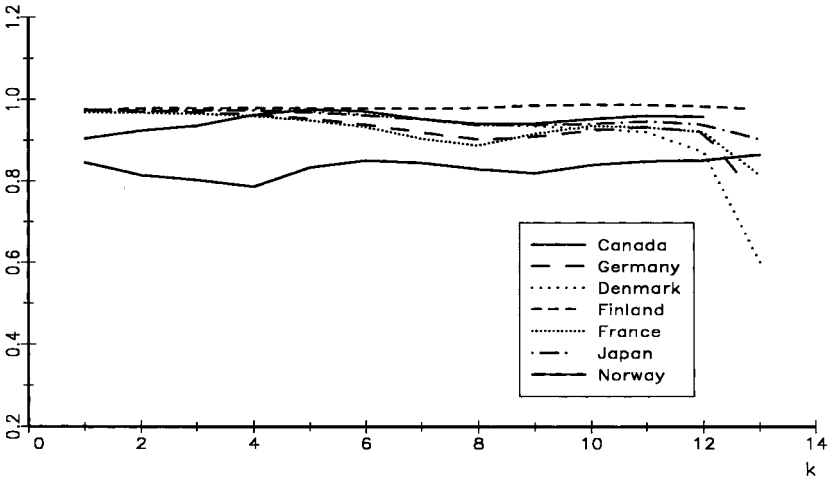

FIG. 4.—MSE decomposition: output prices. $\mathrm{MSE}(x_{t+k} - x_t)/[\mathrm{MSE}(x_{t+k} - x_t) + \mathrm{MSE}(y_{t+k} - y_t)]$, 1973–89.

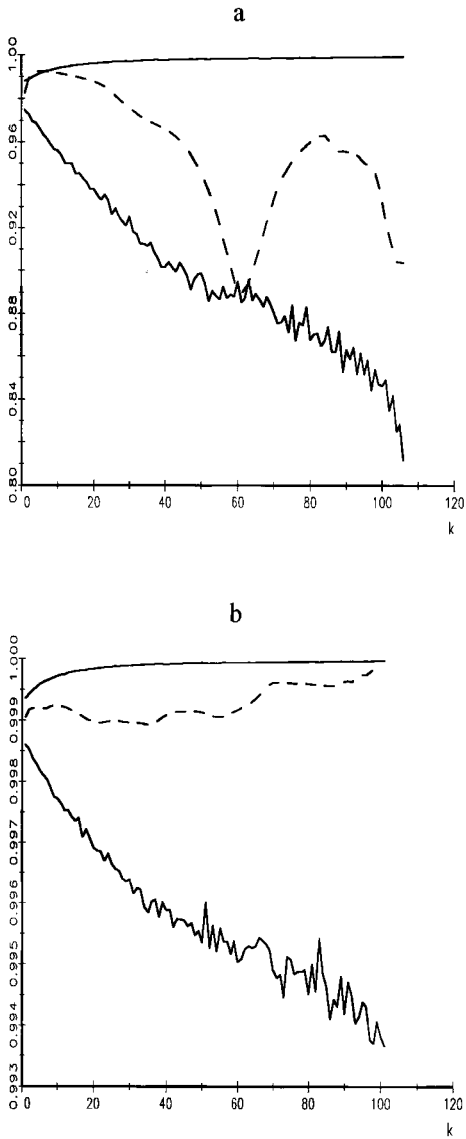

FIG. 5.—MSE decomposition: personal consumption expenditure prices. MSE ratio estimates and 95 percent confidence intervals: *a*, Canadian/U.S. consumption deflator real exchange rate. *b*, Japanese/U.S. consumption deflator real exchange rate. *c*, France, Italy, Sweden, and United Kingdom.

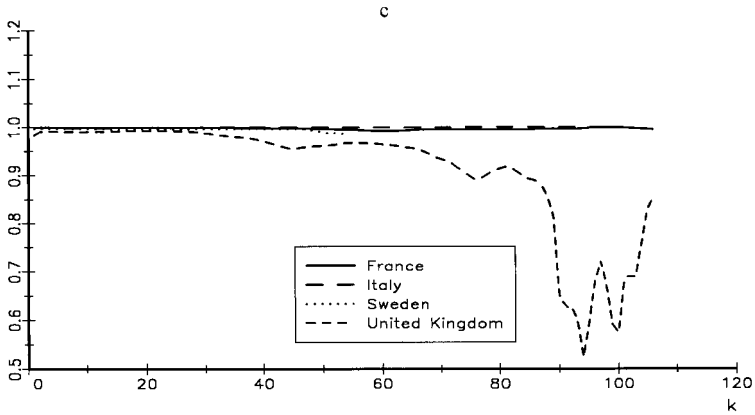

FIG. 5.—(*Continued*)

zons, the traded-goods component accounts for nearly all of the movement in the real exchange rate. For the U.S.-U.K. real exchange rate, x_t is responsible for over 95 percent of the movement out to 15 years. The fraction tails off to around 85 percent at the 20-year horizon.

As in the previous two sections, I plot 95 percent confidence bands for the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component for Canada (fig. 5*a*) and Japan (fig. 5*b*). Again, these bands are computed under the null that $x_{t+1} - x_t$ and $y_{t+1} - y_t$ are independent, normally distributed random variables with means and variances equal to their sample means and variances. I find that for the two real exchange rates plotted, as well as for all the other real exchange rates (whose confidence bands are not plotted because of space considerations), the sample value of the ratio of the MSE of the traded-goods component to the MSE of the real exchange rate lies within the 95 percent confidence interval at all horizons.

IV. Producer Price Index

This section uses the overall PPI as an index of traded-goods prices. The traded-goods component, x_t, is constructed as

$$x_t = s_t + \ln(\mathrm{PPI}_t^*) - \ln(\mathrm{PPI}_t). \tag{2}$$

The nontraded component, y_t, is calculated as

$$y_t = \ln(\mathrm{CPI}_t^*) - \ln(\mathrm{PPI}_t^*) - [\ln(\mathrm{CPI}_t) - \ln(\mathrm{PPI}_t)]. \tag{3}$$

The data are monthly from January 1972 to late 1997 for 16 coun-

tries: Austria, Belgium, Canada, Denmark, Finland, Germany, Greece, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Switzerland, the United Kingdom, and the United States.[10]

This data set has the advantage that it covers many more countries than any of the others. However, there are at least four problems with these data that seriously damage their worth.

1. Using the aggregate PPI as a measure of traded-goods prices is crude. Clearly, some output is nontraded. Indeed, in Section II, I use only producer price data (from a different source) and classify some sectors as tradable and some as nontradable.

2. In contrast to the other data, the measures of traded-goods prices and nontraded-goods prices come from different surveys. Prices for similar goods might be measured in different ways in the two indexes. The PPI and CPI measures, for example, may have different methods of averaging recordings of disparate prices for the same good; they may survey different locations; they may adjust for changes in quality differently, and so forth. There is no attempt to reconcile these differences, as there would be when prices all come from the same source.

3. Equation (3) allows us to construct an accurate measure of y_t only if the aggregate price index is a geometric average of traded-goods prices and nontraded-goods prices. However, the CPI (which we are using as the aggregate price index) is not constructed this way. So even if the PPI were a good measure of traded-goods prices, equation (3) would not give us a good measure of y_t.

4. In the previous sections, the decomposition of q_t was uncomplicated by worries of how to treat comovements of x_t and y_t because these two variables were nearly uncorrelated. But for the measures in this section, x_t and y_t are negatively correlated. The reason is probably that $\ln(\mathrm{PPI}_t^*) - \ln(\mathrm{PPI}_t)$ appears with a positive sign in x_t and a negative sign in y_t. The decomposition of q_t depends on how comovements of x_t and y_t are handled in this case.

Figure 6 displays the decomposition of the MSE given by equation (B2) from Appendix B. Note that it is possible for x_t to be responsible for more than 100 percent of the MSE of q_t under this formula if the comovements of x_t and y_t are sufficiently negative. Formula (B2) arbitrarily classifies half of the comovements as being caused by changes in x_t and half as coming from changes in y_t.

The general picture in figure 6 is that the x_t component accounts for over 70 percent of real exchange rate movements out to horizons of 10 years. The Canadian and Belgian real exchange rates are minor exceptions because at some horizons less than 10 years the x_t compo-

[10] Portuguese and Italian data end in 1986 and 1985, respectively.

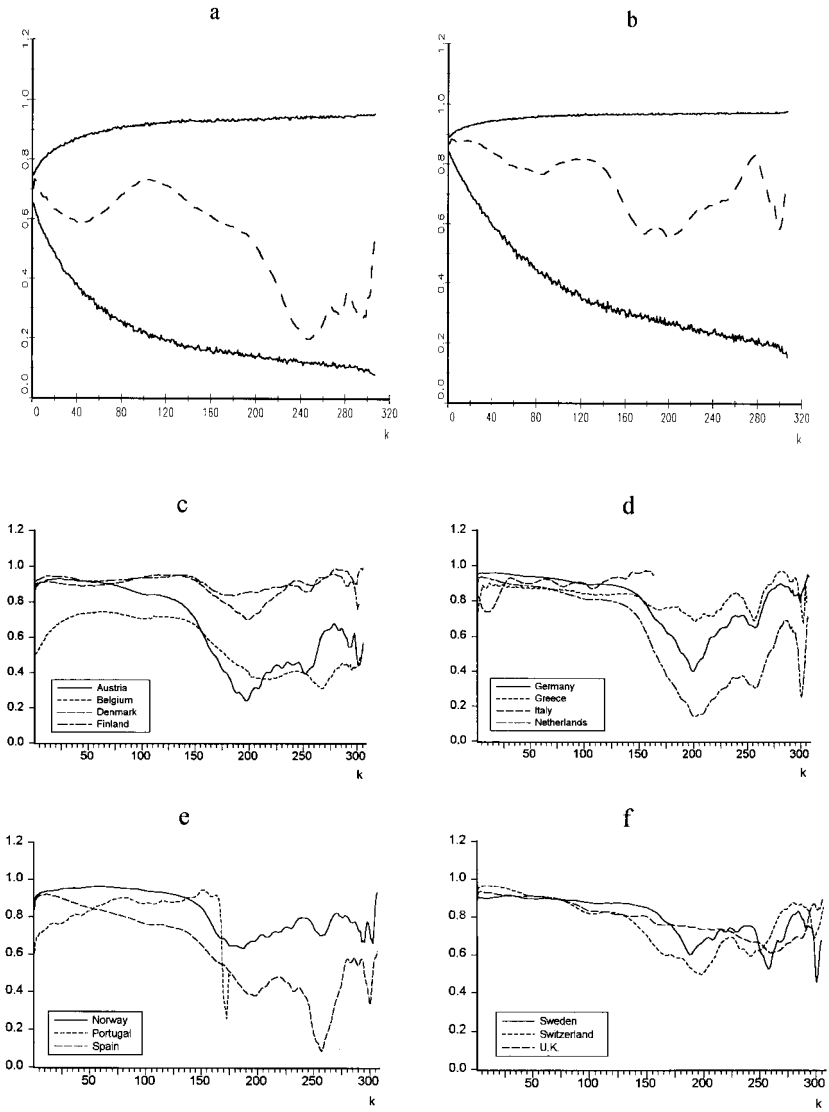

FIG. 6.—MSE ratio estimates and 95 percent confidence intervals. PPI is the price of traded goods, and CPI is the price of all goods. *a*, Canadian/U.S. real exchange rate, January 1972 to November 1997. *b*, Japanese/U.S. real exchange rate, January 1972 to November 1997. *c*, Austria, Belgium, Denmark, and Finland, January 1972 to December 1997. *d*, Germany, Greece, Italy, and the Netherlands, January 1972 to December 1997. *e*, Norway, Portugal, and Spain, January 1972 to December 1997. *f*, Sweden, Switzerland, and the United Kingdom, January 1972 to December 1997.

nent is responsible for only around 60 percent of real exchange rate movements. For most of the countries, x_t determines greater than 80 percent and in many cases greater than 90 percent of movements in q_t at horizons less than 10 years.

Figures 6*a* and 6*b* plot 95 percent confidence intervals for the U.S.-Canadian and U.S.-Japanese decompositions. As in the previous sections, the sample value of the ratio of the MSE of the traded-goods component to the MSE of the real exchange rate lies within the 95 percent confidence interval at all horizons. These confidence intervals are calculated under the null hypothesis that $x_{t+1} - x_t$ and $y_{t+1} - y_t$ are independent, normally distributed random variables with means and variances equal to their sample means and variances. In contrast to the previous sections, the null is rejected at some horizons (particularly longer ones) for some currencies because, as I have already noted, x_t and y_t are negatively correlated. The decompositions depicted in figures 6*a*–6*f* will depend on how comovements are treated for those countries.

V. The U.S.-Japanese Real Exchange Rate

The nontraded-goods price component, y_t, accounts for little movement in the real exchange rate at any of the horizons for which one can reliably measure the MSE. Here we shall examine in more detail the real exchange rate between the United States and Japan.

Figure 7 plots real CPI and PPI exchange rate indexes with

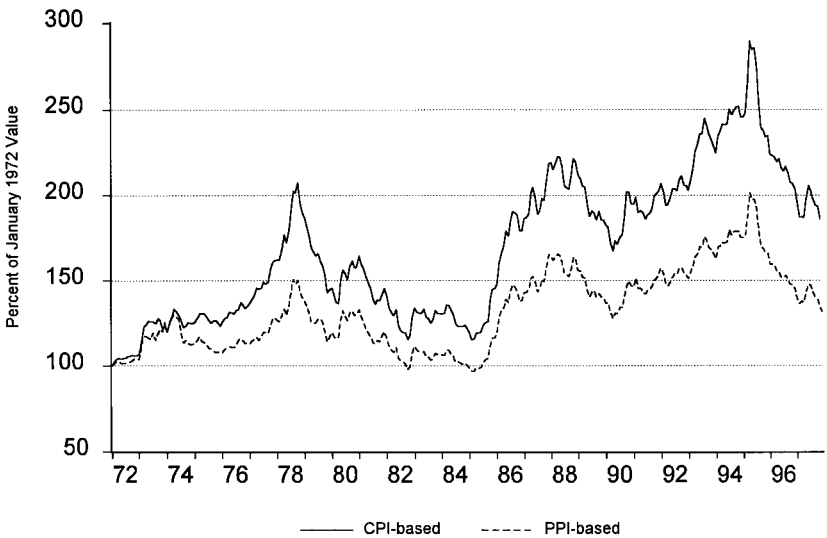

FIG. 7.—U.S./Japanese real exchange rates: CPI-based and PPI-based measures

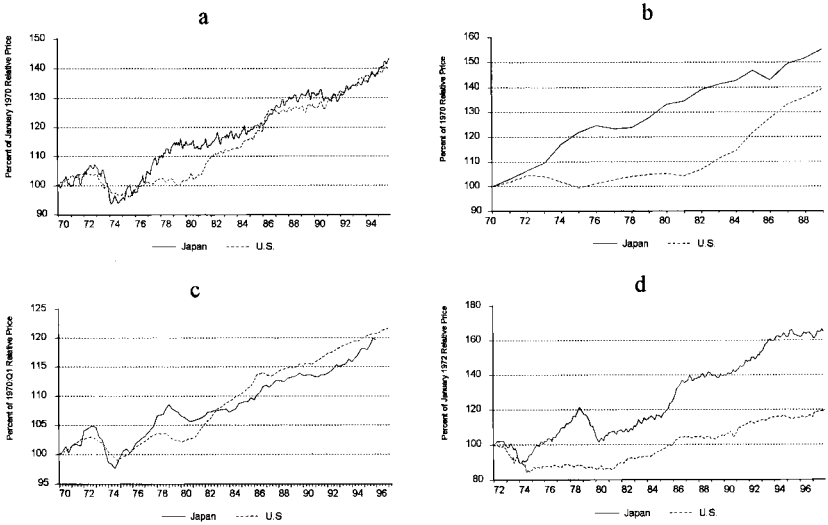

FIG. 8.—Price of nontraded goods relative to traded goods in Japan and the United States. *a*, Consumer prices in Japan and the United States. *b*, Output prices in Japan and the United States. *c*, Expenditure deflator price in Japan and the United States. *d*, Ratio of CPI to PPI in Japan and the United States.

monthly data from January 1972 to November 1997. For the CPI exchange rate, the dollar has depreciated in real terms by about 90 percent, whereas in PPI terms the real depreciation has been about 40 percent. There has also been dramatic short-term and medium-term volatility. Between 1977 and 1983, the dollar gained about 80 points on the CPI real exchange rate index and about 50 on the PPI real exchange rate index. Large gains for the dollar were also recorded in the 1988–90 and post-1995 periods.

Figure 8 displays plots of the price of nontraded goods relative to traded goods in Japan and the United States for the four measures of prices that were used in Sections I–IV.

Consider figure 8*a*, for consumer prices. Nontraded-goods prices have risen steadily relative to traded-goods prices in Japan since 1970. This is consistent with the nontraded-goods explanation for the real appreciation of the yen. From equation (1), as relative nontraded-goods prices increase in Japan, the real yen rises. But there are four reasons why this movement is not responsible for much of the movement in the real exchange rate: (1) The increase in the relative price of nontraded goods in Japan was about 40 percent during this period, but the appreciation of the CPI real exchange rate was 90 percent. (2) There are no large downswings in the relative price of nontraded goods in Japan, but there were periods of

dramatic real depreciation of the yen. (3) In equation (1), the effect of nontraded-goods prices is diminished in the real exchange rate calculation because it is multiplied by the share of nontraded goods in Japanese consumption, which is calculated to be about .31. (4) The most important reason why the nontraded-goods component, y_t, can explain so little of the real exchange rate movement is that it subtracts off $\alpha(p_t^N - p_t^T)$ from $\beta(p_t^{N*} - p_t^{T*})$. It is the relative relative price of nontraded goods that matters. As figure 8 shows, the relative price of nontradables in the United States has closely mirrored the relative price of nontradables in Japan. Their movements cancel each other.

The graphs for the relative prices of nontraded goods constructed from the OECD output prices (fig. 8*b*) and for the deflator for personal consumption expenditures (fig. 8*c*) give a similar perspective. Movements of nontraded-goods prices in Japan are generally offset by movements of nontraded-goods prices in the United States.

Figure 8*d*, which plots the ratio of the CPI to the PPI in each country, is a little different. In Japan, since 1972, the CPI has risen by 60 percent relative to the PPI, whereas in the United States the increase in the CPI/PPI ratio has been less than 20 percent. The movement in these prices is still not enough to account for much of the real exchange rate movement (as fig. 6*b* documents), but it does appear to give more credence to the nontraded-goods price avenue than any of the other measures.

One potential problem with the measures of traded-goods prices is that there are extensive producer services—distribution, retailing, advertising, and so forth—whose costs are included in the traded-goods price. Since these marketing services are largely nontradable, this hypothesis could revive the possibility that the relative price of nontradables can account for much of the real exchange rate movement. It is possible that the rapid increase in the price of marketing services in Japan accounts for the real appreciation of the yen.

We have already seen one bit of evidence that weighs in against this hypothesis. Even when we examine the OECD output prices in Section II, we find that the traded-goods component accounts for almost all the real exchange rate movements. The marketing and distribution prices should not be an important component of these output prices.

More direct evidence is reported in figure 9, which graphs monthly prices of marketing services in Japan relative to the CPI for all goods beginning in January 1985. The chart shows prices for all services, finance and insurance, real estate services, transportation services, leasing and renting, and information services. The data are

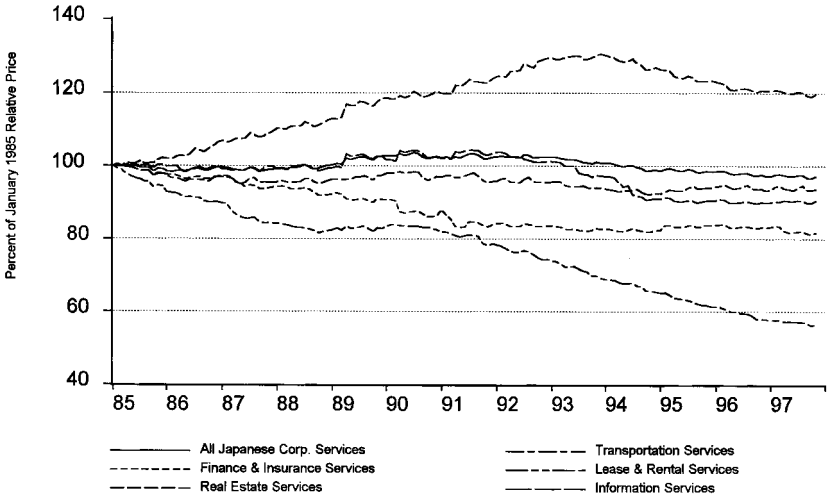

FIG. 9.—Price of Japanese corporate services relative to the overall Japanese CPI

available for another 30 categories of corporate services, but they are not graphed because of space considerations.

If the prices of these services are to explain the real appreciation of the yen and the large swings in the yen's real value, they must rise relative to the overall CPI during this period and experience wide swings. In fact, most of these prices fall relative to the CPI. Only six of the 36 noticeably increase against the CPI: real estate services, other services, building maintenance, civil engineering and construction, judicial and accounting services, and industrial waste disposal. The index for all corporate services holds steady relative to the CPI over the entire period, as can be seen in figure 9. So the prices of marketing services do not move significantly and do not appear to be able to account for the real exchange rate movements.

VI. Bias in Traded-Goods Measurement

At several points I have discussed the possibility that the traded-goods prices are mismeasured because they may contain a nontraded component. The previous section shows that the prices of marketing services in Japan do not show any dramatic movement that is comparable to swings in the real exchange rate. But the MSE of the price of nontraded to traded goods could be significantly understated even if there were no dramatic movements in the price of the nontraded marketing services.

Suppose, in fact, that nontraded marketing services are perfectly correlated with the measure of other nontraded-goods prices, p_t^N. But the measure of traded-goods prices, p_t^T, is actually a weighted average of the nontraded-goods price and the price of goods that are truly traded, p_t^C:

$$p_t^T = \gamma p_t^N + (1 - \gamma) p_t^C.$$

Likewise, in the foreign country, γ is the weight that nontraded marketing services take in the measure of traded-goods prices.

In this case, the weight that nontraded goods receive in the overall price index domestically is actually $\alpha + \gamma(1 - \alpha)$, whereas I have taken it to be just α. The true measure of y_t that one would like is

$$\hat{y}_t = [\beta + \gamma(1 - \beta)](p_t^{N*} - p_t^{C*}) - [\alpha + \gamma(1 - \alpha)](p_t^N - p_t^C).$$

Since I actually use y_t, there is bias in the measure of $\hat{y}_t$. We can write y_t as

$$y_t = \beta(1 - \gamma)(p_t^{N*} - p_t^{C*}) - \alpha(1 - \gamma)(p_t^N - p_t^C).$$

While the domestic relative price of nontraded goods should be getting a weight of $\alpha + \gamma(1 - \alpha)$ as in the expression for $\hat{y}_t$ (and $\beta + \gamma(1 - \beta)$ for foreign goods), we are actually giving them a weight of $\alpha(1 - \gamma)$ (and $\beta(1 - \gamma)$ for foreign goods) when we use y_t as our measure. This biases downward the importance of the relative price of nontraded goods. When γ is large, the bias could be severe.

The problem is that we do not measure p_t^C and p_t^{C*} directly. We shall, however, construct measures by using the relation

$$p_t^C = \frac{1}{1 - \gamma}(p_t^T - \gamma p_t^N). \tag{4}$$

The value of γ is unknown, but we shall experiment with different values. We can use this measure of p_t^C (and the analogous measure for p_t^{C*}) to get a measure of $\hat{y}_t$.

Before proceeding in this direction, note that this calculation generates a series for $\hat{y}_t$ that is too volatile. We have assumed that the price of marketing services is perfectly correlated with p_t^N. Suppose instead that p_t^N measures the price of marketing services, p_t^{MS}, with measurement error u_t:

$$p_t^{MS} = p_t^N + u_t,$$

and likewise for the foreign country. Ideally, $\hat{y}_t$ would be measured as a weighted average of the relative prices of p_t^N and p_t^{MS} to p_t^C:

$$\beta(p_t^{N*} - p_t^{C*}) + \gamma(1 - \beta)(p_t^{MS*} - p_t^{C*}) - \alpha(p_t^N - p_t^C) - \gamma(1 - \alpha)(p_t^{MS} - p_t^C).$$

But when measurement error is introduced into p_t^{MS}, our measure of $\hat{y}_t$ is actually

$$\beta(p_t^{N*} - p_t^{C*}) + \gamma(1 - \beta)(p_t^{MS*} - p_t^{C*}) - \alpha(p_t^N - p_t^C) - \gamma(1 - \alpha)(p_t^{MS} - p_t^C) + \frac{\gamma}{1 - \gamma}(u_t - u_t^*).$$

This introduces volatility into our measure of $\hat{y}_t$ that can be substantial when γ is large.

With this caveat in mind, note that using equation (4) (and its equivalent for p_t^{C*}), we can write $\hat{y}_t$ in terms of variables that are measurable:

$$\hat{y}_t = \left(\beta + \frac{\gamma}{1 - \gamma}\right)(p_t^{N*} - p_t^{T*}) - \left(\alpha + \frac{\gamma}{1 - \gamma}\right)(p_t^N - p_t^T).$$

So we can take our previous measure of y_t and simply increase the weights from β and α to $\beta + [\gamma/(1 - \gamma)]$ and $\alpha + [\gamma/(1 - \gamma)]$.

Figure 10 reproduces the MSE decomposition using CPIs over the floating exchange rate period that was presented in figure 2. However, the weights on the relative prices of nontraded goods have been increased. In figure 10*a*, γ is set to 1/3 (which means that 0.5 is added to β and α).[11] This choice of γ is arbitrary. Its interpretation is that the share of marketing services in the final-goods prices is 1/3. We see that the fraction of the MSE attributable to the traded-goods component, x_t, is lower than in figure 2. However, it is still very large: x_t accounts for over 90 percent at all but the longest horizons except for the U.S.-Canadian real exchange rate.

Figure 10*b* uses a value of γ equal to 2/3 (which means that 2.0 is added to β and α). Both because of the measurement problems introduced into our measure of $\hat{y}_t$ that overstate its volatility and because the marketing share is unlikely to be as large as 2/3, the numbers presented in figure 10*b* are probably a lower bound on the share of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component, $x_{t+k} - x_t$. Even so, the traded-goods component is responsible for a large share, especially at short to medium horizons.

I conclude that even when one takes into account the possibility

[11] The term $s + p_t^{T*} - p_t^T$ is used in constructing x_t rather than $s + p_t^{C*} - p_t^C$. Because $p_t^{T*} - p_t^T$ is less volatile than $p_t^{C*} - p_t^C$, fig. 10 biases down the contribution of the x_t component.

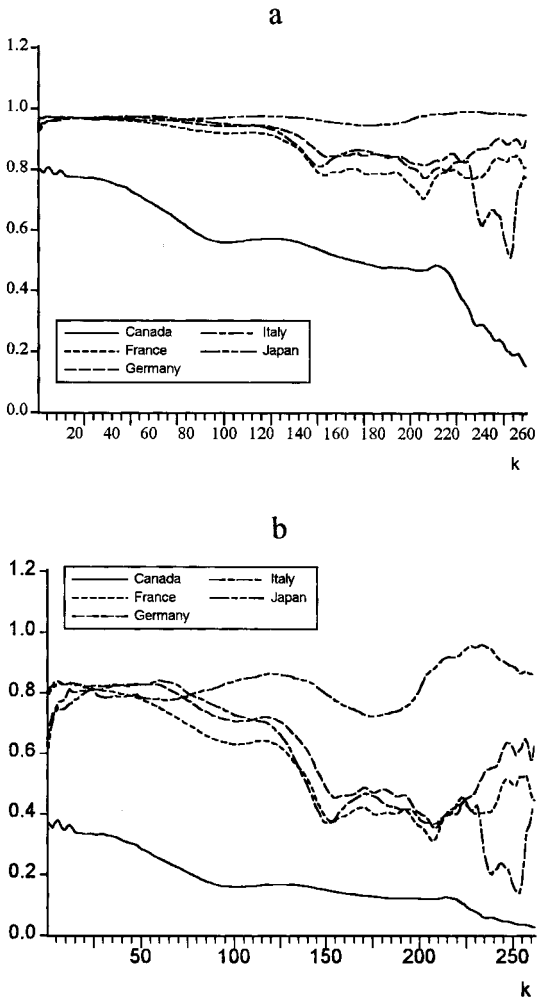

FIG. 10.—MSE decomposition: measurement error in traded-goods prices, March 1973 to December 1995. *a*, $\gamma = 1/3$. *b*, $\gamma = 2/3$.

that there is a nontraded component in the measure of traded-goods prices, the relative price of nontraded goods contributes little to the movements in real exchange rates.

VII. Conclusions

So far no explanation has been offered for why the movements in the relative price of traded goods are so important for the real ex-

change rate. One possibility is terms-of-trade movements. If there are fluctuations in the relative prices of goods that constitute the traded-goods price indexes and if the weights these goods receive are different in the U.S. and foreign price indexes, then the relative traded-goods price indexes will fluctuate.

It is difficult to investigate this avenue with my data since one cannot disaggregate the traded-goods price indexes much. With the consumer price data, the indexes are composed from only two subindexes: for food and for all goods not including food. For the OECD price data, the traded-goods price index is composed of subindexes for 12 categories of goods and services. None of the other traded-goods price indexes can be disaggregated.

Appendix C reports the outcome of investigating the terms-of-trade hypothesis for the CPI and OECD output price indexes. There, I find that this hypothesis explains very little because the weights of the subcategories in the overall price indexes are not too different across countries. So even if there is a lot of volatility in relative price movements among the subcategories, it does not have much of an effect on the overall relative price.

It should be noted that all the real exchange rates examined in this study pertain to high-income countries. As table 3 of Rogoff (1996) demonstrates, the behavior of real exchange rates for low-income countries might be very different. The absolute purchasing power of the currencies of low-income countries is much higher than for high-income countries, although within the two groups there seems to be little relationship between income and purchasing power. It seems likely that the disparity between low-income and high-income countries' purchasing powers reflects differences in the relative prices of nontraded goods.

It is tempting to attribute the movements in the traded-goods price indexes to failures of the law of one price. That conclusion would be consistent with the findings of a number of recent studies (Engel 1993; Froot et al. 1995; Rogers and Jenkins 1995; Engel and Rogers 1996, 1998; Knetter 1997). However, there are many issues to be resolved in that area: What determines export prices to different regions from the same country? How important is pricing to market in determining real exchange rate movements? What systematic relationship is there between the price of a good at the port and at the consumer outlet? To what extent do the degree of product differentiation and the competitiveness of the industry determine differences in prices between locations? What role does nominal price stickiness play, and how is it related to these other questions? The key to understanding real exchange rate movements may be found

in the answers to these questions. But the evidence of this paper suggests that the relative price of nontraded goods has little import for understanding U.S. real exchange rate movements over the short and medium run.

Finally, it should be emphasized that this study cannot shed light on the importance of the relative price of nontraded goods in determining the real exchange rate in the long run. Engel (1996), in fact, argues that nontraded-goods prices could have a large impact on real exchange rates at long horizons. In that study, a simple model for the U.S.-U.K. real exchange rate is calibrated (using the data from Sec. III of this paper) in which the y_t component is nonstationary and x_t is stationary. At horizons of 100 years, the y_t component accounts for about half of the variance of the real exchange rate. Yet, using Monte Carlo exercises, I find that one is very likely to conclude that the real exchange rate is stationary. That is, even in 100 years of data I cannot detect the nonstationary y_t component. Inference about the presence of a unit root component is difficult because the stationary x_t variable has such a high variance that it masks the movements in the nonstationary y_t variable. Certainly in the 20–30-year samples used in this paper, one would not be able to discern whether nontraded-goods prices had a significant influence on real exchange rate movements in the very long run.

Appendix A

Data

Nominal exchange rates are taken from Datastream. Period average exchange rates from the International Monetary Fund's International Financial Statistics database are used for 1972–97. In general, monthly price indexes are constructed from averages of prices observed throughout the month. So monthly average exchange rates are appropriate: rates measured at a point in time would overstate the volatility of exchange rates compared to prices. For 1969–71, exchange rates are period averages from Datastream's national government series. End-of-period exchange rates from the OECD database are used for 1960–69. Since exchange rates are generally fixed in this period, there is little difference between period averages and end-of-period rates.

Seasonally unadjusted monthly data from January 1972 to December 1995 on consumer prices are used in Section I. The data are taken from Datastream's OECD database on all items (AI), all goods less food ($AGLF$), food (F), services less rent (SLR), and rent (R). The weights in the price index are constructed from the regression

$$\Delta(ai - r) = \phi_1\Delta(aglf - r) + \phi_2\Delta(f - r) + \phi_3\Delta(slr - r) + \epsilon,$$

where Δ is the first-difference operator, and lowercase letters are used to denote natural logarithms. Then I construct

$$p^T = \left(\frac{\phi_1}{\phi_1 + \phi_2}\right) \cdot aglf + \left(\frac{\phi_2}{\phi_1 + \phi_2}\right) \cdot f$$

and

$$p^N = \left(\frac{\phi_3}{1 - \phi_1 - \phi_2}\right) \cdot slr + \left(\frac{1 - \phi_1 - \phi_2 - \phi_3}{1 - \phi_1 - \phi_2}\right) \cdot r.$$

Section II constructs price indexes for each of 19 sectors as the ratio of nominal to real output, measured annually, taken from the OECD's international sectoral database. The nontraded- and traded-goods price indexes are weighted averages of the price indexes of each sector; the weights are fixed and taken as the average share of each sector's output in total output. The data ranges are Canada, 1961–89; Germany, 1960–90; Denmark, 1966–90; Finland, 1960–90; France, 1970–90; Japan, 1970–90; Norway, 1962–88; and United States, 1960–89.

The prices in Section III are personal consumption deflators from Datastream's OECD quarterly accounts. The data ranges for Canada, France, and the United Kingdom are 1970:1–1997:1; for the United States, 1970:1–1997:2; for Japan, 1970:1–1996:1; for Italy, 1970:1–1995:3; and for Sweden, 1980:1–1994:2. Price indexes are computed as the ratio of nominal to real expenditures, with commodities classified as traded goods and services as nontraded. The data for Canada, France, Italy, and the United States were seasonally adjusted. For Japan and Sweden they were unadjusted, but were adjusted using a multiplicative seasonal adjustment.

In contrast to all the other price data used in this study, these prices are seasonally adjusted. This is likely to make the traded-goods component, x_t, relatively more important in explaining short-term movements in q_t. Seasonal adjustment tends to smooth the price series over short intervals. But x_t also contains the nominal exchange rate, which is not seasonally adjusted, so x_t will be relatively more volatile.

Section IV uses monthly, not seasonally adjusted, producer and consumer price indexes from Datastream's International Monetary Fund series with four exceptions. The Belgian, German, and Norwegian producer price series and the Italian consumer price series were taken from Datastream's national account data. All data begin in January 1972. The data for Germany and the United States end in 1997:12; Canada, Japan, the Netherlands, Norway, Spain, and Switzerland, in 1997:11; the United Kingdom, 1997:10; Sweden, 1997:9; Finland, 1997:7; Portugal, 1986:12; and Italy, 1985:12.

The prices of corporate services used in Section V are taken from Datastream's national government series. The data are monthly from January 1985 to November 1997, not seasonally adjusted.

Nontraded goods' weight in the price indexes of Sections I–III are shown in table A1.

TABLE A1

Country	Section I	Section II	Section III
Canada	.3942	.6110	.3147
Denmark	...	.6524	...
Finland	...	.5506	...
France	.2530	.6261	.2805
Germany	.2811	.5254	...
Italy	.2436	...	.2484
Japan	.3114	.5919	.3292
Norway	...	.5561	...
Sweden	...	...	.2837
United Kingdom	...	...	.4024
United States	.4590	.6450	.3379

Appendix B

Decomposition of Real Exchange Rate Movements

I employ two different methods for measuring the fraction of the MSE of $q_{t+k} - q_t$ accounted for by the MSE of the traded-goods component, $x_{t+k} - x_t$. The issue is how to deal with comovements of x_t and y_t. The first decomposition is

$$\frac{\mathrm{MSE}(x_t - x_{t-n})}{\mathrm{MSE}(x_t - x_{t-n}) + \mathrm{MSE}(y_t - y_{t-n})}, \tag{B1}$$

and the second is

$$\frac{\mathrm{MSE}(x_t - x_{t-n}) + \mathrm{mean}(x_t - x_{t-n})\,\mathrm{mean}(y_t - y_{t-n}) + \mathrm{cov}(x_t - x_{t-n}, y_t - y_{t-n})}{\mathrm{MSE}(q_t - q_{t-n})}. \tag{B2}$$

The mean-squared error is defined as

$$\mathrm{MSE}(x_t - x_{t-n}) = \mathrm{var}(x_t - x_{t-n}) + [\mathrm{mean}(x_t - x_{t-n})]^2.$$

Measure (B1) ignores comovements. Measure (B2) attributes half of the comovements to x_t. In Sections I–III, the two measures give very similar results because comovements of x_t and y_t are so small. In fact, I cannot reject that these two components are independent random walks for all the prices in Sections I–III. As the text notes, comovements are more important for the prices used in Section IV. That section uses (B2) for the decomposition.

The mean change is calculated as

$$\mathrm{mean}(x_t - x_{t-n}) = \frac{n}{N-1}(q_N - q_1),$$

where N is the sample size.

The variance calculation makes the small-sample correction suggested by Cochrane (1988):

$$\mathrm{var}(x_t - x_{t-n}) = \frac{N}{(N-n-1)(N-n)} \sum_{j=1}^{N-n} [x_{j+n} - x_j - \mathrm{mean}\,(x_{j+n} - x_j)]^2.$$

A small-sample correction would not be needed for a variance decomposition since the bias correction would cancel out in the numerator and denominator of the decomposition. But the MSE decomposition would overemphasize the squared drift if a downward-biased measure of the variance were used.

To construct confidence intervals, I first calculate the sample mean and variance of $x_{t+1} - x_t$ and $y_{t+1} - y_t$. Then 5,000 artificial series of length $N - 1$ of $x_{t+1} - x_t$ and $y_{t+1} - y_t$ are created using the normal random number generator (*rndn*) on Gauss with mean and variance equal to those of the data. These are cumulated to create artificial series of x_t and y_t, which are now, by construction, independent random walks. Then the nth differences of these series are taken, and the MSE decompositions recorded. The confidence intervals exclude the largest 125 and smallest 125 ratios that are calculated. This process is repeated to get the confidence interval at each of the n horizons. (New artificial series are created for the calculations at each horizon because the computer program would be inefficient if it had to save all 5,000 series to do the calculations at each horizon.)

Appendix C

Terms-of-Trade Movements

This Appendix briefly addresses the issue of why movements in the relative price of traded goods are so large. One possibility is that the law of one price fails. Another possibility is that the traded-goods price indexes weight goods differently in the two countries. Suppose that

$$p_t^T = \nu p_t^1 + (1 - \nu) p_t^2$$

and

$$p_t^{T*} = \eta p_t^{1*} + (1 - \eta) p_t^{2*}.$$

Even if the law of one price held for each good, so $p_t^j = s_t + p_t^{j*}$ ($j = 1, 2$), if $\nu \neq \eta$, then $p_t^T - s_t - p_t^{T*}$ will change as p_t^1 moves relative to p_t^2.

To gauge the importance of this possibility, I recalculate the foreign traded-goods price index using U.S. weights to see whether this significantly lowers the MSE of $p_t^T - s_t - p_t^{T*}$. If terms-of-trade movements were important, the MSE should decline under this recalculation. In fact, the MSE (and the squared drift and variance) does not decline. The same is true for the opposite tack of calculating the U.S. price index using foreign weights.

For example, figure C1 plots the ratio of the MSE of $p_t^T - s_t - p_t^{T*}$ when foreign weights are used in the U.S. price index to the MSE when the U.S. price index uses its own weights. Figure C1*a* shows this ratio using the consumer prices of Section I. In all cases, the ratio is close to unity: there is little effect from using foreign weights. Figure C1*b* plots the ratio using the output prices from Section II. Again, all the ratios are near one.

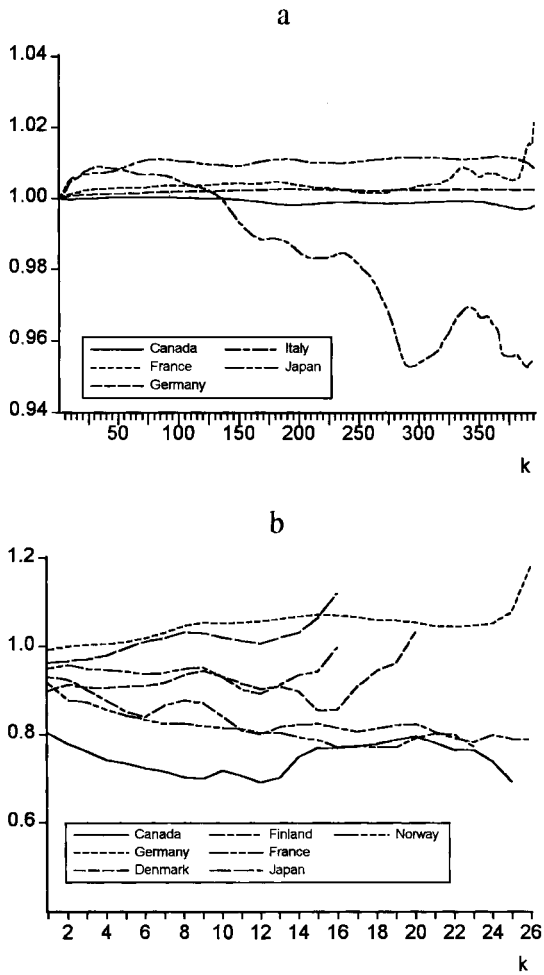

FIG. C1.—MSE effect of variation in the terms of trade. *a*, CPI prices, January 1962 to December 1995. *b*, Output prices, 1960–89.

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