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In This Issue . . .

Recent theoretical explanations and empirical analyses of exchange rates emphasize the role of asset markets rather than trade flows. Many argue that forward exchange rates and the future spot exchange rates they may predict are primarily determined by interest and inflation rate differentials between countries. In the first article in this *Review*, "Forward Exchange Rates in Efficient Markets: The Effects of News and Changes in Monetary Policy Regimes," Mack Ott and Paul T.W.M. Veugelers investigate the extent to which errors in forward exchange rate predictions of future spot exchange rates have been influenced, on the one hand, by changes in interest and inflation rates and, on the other, by changes in the policy stance of the U.S. monetary authority.

The authors find that changes in interest differentials explain a portion of forward rate forecast errors, especially during the period of U.S. monetary aggregate targeting, October 1979 to September 1982, and that changes in the U.S. monetary policy regime alter the risk premium in forward exchange rates. The significant divergencies between the forward and spot exchange rate relations under different U.S. monetary policy regimes suggest that credible goals for monetary policy may be as important as the mechanical details of that policy's execution.

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In recent years, federal payments to farmers for both loans and purchases of farm products have set new records. In the second article in this issue, "How Federal Farm Spending Distorts Measures of Economic Activity," John A. Tatom explains how transactions by the Commodity Credit Corporation (CCC) are treated in the National Income and Product Accounts (NIPA).

Tatom shows that the volatile, quarter-to-quarter pattern of CCC payments to farmers affects measures of farm, business, government and overall economic activity. According to the author, the recent unusual developments have complicated the interpretation of some key measures of economic performance. He points out that adjusting for these movements can alter significantly conclusions about the short-term performance and economic outlook for federal purchases, business inventory investment and final sales in the economy.

Tatom explains that analysts are likely to be misled about the economy's short-run economic performance unless they properly adjust the NIPA measures when large changes in CCC purchases occur.

Forward Exchange Rates in Efficient Markets: The Effects of News and Changes in Monetary Policy Regimes

Mack Ott and Paul T. W. M. Veugelers

SINCE the late 1970s, theoretical explanations of exchange rate determination have emphasized the asset approach rather than the expenditure approach.[1] Most of the empirical research applying the asset models of exchange rate determination also subsume the efficient market hypothesis. In this article, we test three efficient market hypotheses bearing on forward exchange rates: First, are forward rates unbiased forecasts of future spot exchange rates? Second, does "news" — in particular unanticipated changes in nominal or real interest differentials — explain forward rate forecast errors? Third, are forward rate forecast errors affected by change in the U.S. monetary policy regime? These hypotheses are tested by examining the forecast errors (the difference between the forward rate and the subsequently observed spot rate) for the U.S. dollar forward rate against the currencies of eight industrialized countries over the latest floating-rate era (1973–85).

EFFICIENT MARKETS AND FORWARD EXCHANGE RATES

The forward exchange rate in an efficient market reflects all the information possessed by individuals active in that market. Thus, in an open market, the forward rate should be an unbiased predictor of the future spot rate.[2] Hence, a regression of the observed spot rate at time t on the forward rate at time t − 1 (where exchange rates are measured by natural logarithms of the dollar prices of foreign exchange),

(1) $s_t = a + b f_{t-1} + e_t$,

should result in an estimated constant not significantly different from zero, an estimated coefficient on

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[1]One rationale for this shift is the observation that the interest rate parity (IRP) postulate of the asset view has held up substantially better than the purchasing power parity (PPP) postulate of the expenditure view; see Mussa (1979) and Frenkel (1981b). The former refers to the equality of asset yields across currencies, while the latter refers to the equality of purchasing power across currencies. PPP frequently, and for protracted periods, has been violated by exchange rates; see Frenkel (1981b). Thus, analysts have been faced with either modifying the PPP assumption and diluting its relevance, or accepting the evidence and developing theories to explain it. Indeed, some authors, Bomhoff and Korteweg (1983) and Darby (1981), argue that changing real exchange rates vitiate the relevance of PPP.

[2]See Dornbusch (1976), Mussa (1979), Frenkel (1981a), Bomhoff and Korteweg (1983) and Edwards (1983b).

the forward rate not significantly different from 1.0, and serially uncorrelated errors (e_t).[3]

Risk Premium

The empirical finding of a significant intercept has been sufficiently frequent in recent research that it is no longer interpreted as a departure from market efficiency. The question, then, is, what does the significant intercept represent?

The current view is that the intercept represents a return to speculation.[4] For example, if real interest rates on U.S. securities are higher than those on foreign securities, investors will shift their portfolios toward the higher-yielding securities denominated in U.S. currency; if these investors are risk-averse to unforeseen changes in currency values, they can hedge by selling the higher-yielding U.S. currency forward and buying their own currency forward. By IRP, the resulting upward pressure on the forward rate must just offset the higher yield obtained on the U.S. securities.[5] Thus, the forward rate in equation 1, in such cases, would overestimate the future spot rate so that the estimated intercept would be negative. Conversely, a higher rate on non-U.S. securities, by the same logic, would imply a positive intercept.

News of Interest Rate Changes

Frenkel (1981a) argues that changes in expectations between the time that the forward rate prediction is made and the spot rate is observed explain the forward errors. These changes in expectations, which he calls news, are based on information revealed after the forward contracts are made but before the spot rates are realized. Thus, unanticipated changes in interest rate differentials between time $t-1$ and t, — one example of news — explain part of the residual between the forward rate forecast f_{t-1} and the realized spot rate s_t. Incorporating this modification into equation 1 yields

$$(2)\ s_t = a + bf_{t-1} + c[(i - i^*)_t - E_{t-1}(i - i^*)_t] + e'_t,$$

where i is an interest rate of the same term as the forward rate with asterisks indicating non-U.S. variables (interest rates are not in logs). Once again, risk-neutrality and efficient markets would imply an insignificant intercept and a slope coefficient of unity; the sign of the coefficient on the news variable, however, would depend upon whether the rise in the interest differential were due to a relative rise in U.S. inflation — in which case it would be positive — or a relative rise in U.S. real interest rates — in which case it would be negative.[6]

Frenkel's proxy for the expected interest rate differential was obtained from a regression of the interest differential on its own lagged values and the lagged forward exchange rate. Estimating this model over 1973–79 for the pound sterling, deutschemark and franc, he found the intercept to be insignificant and the coefficient of the lagged forward rate not significantly different from one; these findings are consistent with the efficient market hypothesis. Moreover, the coefficients on the news variable — the unanticipated interest rate change — were positive, which he interpreted as primarily reflecting the relatively high and rising U.S. expected inflation rate during this period.

THE ROLE OF NEWS IN THE FORWARD EXCHANGE MARKET

An important insight of the asset-market approach to exchange rate determination is the emphasis on expectations. Asset prices are much more dependent

[3]These propositions about the forward exchange rate have not been supported by recent empirical work. For example, Hansen and Hodrick (1980) find significant evidence of risk premia and explanatory power in lagged errors in both the 1920s and 1970s in one- and three-month forward markets. Baillie, Lippens and McMahon (1983), using a time series model on weekly data reject the hypothesis that the forward rate is an unbiased predictor of the future spot rate in weekly data. Fama (1984) argues that the risk premium explains much of the error in the forward rate's forecasts and finds that the risk premium and expected future spot rate are negatively correlated. Jacobs (1982) argued that the forward rate is an imperfect proxy for the expected rate and constructs a time series proxy for the expected rate. Unlike Fama, however, Jacobs found information in the past variables, that is, information not included in the efficiently constructed forward rate at time $t-1$. Jacobs' emphasis on omitted information is analogous to the decomposition suggested by Frenkel (1981a) and elaborated in Isard (1983) and Edwards (1983a, 1983b). Edwards (1983b) finds that market efficiency is not rejected in three out of four currencies in his study once news is included.

[4]Fama (1984) and Hodrick and Srivastava (1985). Hodrick and Hansen (1983) find that significant premia are both common and time varying. Frenkel (1981a) finds that news explains some of the risk premium while Edwards (1983b) finds that the combination of news and a system estimation technique eliminates the significant intercept.

[5]Investors are concerned about after-tax real rates of return; throughout this article we ignore the possibility that long-run real interest differentials may persist due to different tax rates on interest and investment income. Since our tests are on the effects of unanticipated *changes* in interest differentials, this possibility does not affect our results.

[6]An increase in the expected inflation rate differential implies that, in the future, the dollar price of foreign currency will rise faster, and fewer dollars will be demanded because of their higher holding cost; hence, s_t would rise. An increase in the U.S. real interest rate relative to foreign rates would increase the value of the dollar; hence, s_t would fall.

than current goods prices on the anticipated course of future events. Consequently, the role of news is most aptly captured in the change of expectations, not the error between the expected and realized yield differentials.

By an application of IRP and the efficient forward market hypothesis for foreign exchange, we can obtain an alternative form of the news equation 2 estimated by Frenkel. The alternative model takes the form (see shaded insert on the next page):

$$(3)\quad s_t - f_{t-1} = \alpha + \beta\Delta fp_t + \omega_t.$$

This model has the advantage of using a market-implied interest differential as well as directly embodying the change in expectations rather than the empirically derived, expectation error proxy used by Frenkel.

The Distinction Between Real and Nominal News

Frenkel claimed that the positive coefficient on the interest rate news he found during 1973–79 reflected the relatively high and rising U.S. inflation rate during this period. Since the U.S. inflation rate has fallen both absolutely and relative to other nations in the years since 1979, the estimated coefficient on the expected nominal interest differential should be unstable over the full period 1973–85. One way to deal with this problem is to break the period into smaller units, each of which have uniform relative U.S. inflation rates. We, instead, separate the real and inflation components of the nominal news variable. That is, we will view the change in the nominal interest differential as the sum of a change in the expected real yield differential and the change in the expected inflation differential. These components of the news should have different effects on the forward rate errors.

A rise in the real yield on investments in one country relative to those elsewhere, in the absence of capital restrictions, will cause an immediate appreciation in its exchange rate and result in a negative error in equation 3. Such appreciations are transitory because capital inflows will bring down the initially higher yields, while the concomitant outflows raise the yields elsewhere, until equality of yields is restored.[7] Consequently, the very rise in the relative yield that causes a currency to appreciate also creates the anticipation of its subsequent depreciation as yield differences go to zero.

In contrast, an increase in the expected inflation differential primarily alters the rate of depreciation of the exchange rate by changing its PPP level; a rise in the inflation differential causes the exchange rate to rise faster over time by the amount of the inflation increase. The depreciation of the spot rate also will reflect the perceived increase in the holding costs of the country's currency which reduces the quantity demanded.

Thus, express the nominal news as the sum of its real and inflation components,

$$(4)\quad \Delta fp_t = \Delta(r_t - r_t^*) + \Delta(\pi - \pi_t^*),$$

where r_t = expected real interest rate, and
π_t = expected inflation rate.

Then, substitute the right-hand-side of equation 4 into equation 3, to obtain

$$(5)\quad s_t - f_{t-1} = \alpha + \beta_1\Delta(r_t - r_t^*) + \beta_2\Delta(\pi_t - \pi_t^*) + \varepsilon_t.$$

In equation 5, α is non-zero in the presence of a risk premium, β_1 is negative (since an unanticipated relative rise in U.S. real rates lowers s_t, implying $s_t - f_{t-1} < 0$), β_2 is positive but smaller than β_1 (since a rise in the relative U.S. inflation rate will cause a change in the rate of depreciation of the dollar, and, through decreased demands for transaction balances, some decline in its level), and ε_t is a serially uncorrelated disturbance term.

Another Kind of News: Changes in Monetary Policy Regimes

The estimated parameters of an economic relation reflect the perceived policy stance of the government and monetary authorities. Thus, as Lucas (1974) argued, changes in policy, either broad goals such as the desired inflation rate or narrower ones such as the method in which the policy is implemented, may alter the public's response to prices and other information.[8]

[7]See Dornbusch (1976), Isard (1983), and Edwards (1983a). Nonetheless, the existence of risk premia implies that interest differences have persisted for some time in open capital markets; see Fama (1984). Hodrick and Hansen (1983) find these risk premia to be nonconstant and that their time variation is *not* summarized by nominal interest rate movements.

[8]We abstract from changes in the long-run real exchange rate in this analysis. That is, different rates of capital or human capital investment will cause different rates of productivity growth, or resource price changes that can alter the real exchange rate; see Darby (1980), Bomhoff and Korteweg (1983). Also, a reduction in the security of property rights can make investment in one currency less attractive than investments in other currencies, depreciating the currency and raising its real yields; see Dooley and Isard (1980). An apt application of the Dooley-Isard hypothesis may be the change in the French government in 1981, which was followed by significant nationalizations — especially in the banking sector. In our analysis, the only structural change considered is the U.S. monetary policy regime.

Forward Exchange Rate Errors, Efficient Markets and the News: the Role of the Forward Premium

In its strong form, the efficient market hypothesis implies that the intercept in equation 1 will be zero and the coefficient of the lagged forward rate will be unity. Consequently, the error term, e_t, is simply the error of the forward rate's forecast of the spot rate,

$$(2.1)\quad s_t - f_{t-1} = e_t.$$

Frenkel's insight concerning the role of news is to argue that this error is due to information revealed after $t-1$ (but before t) which alters expectations and, hence, s_t:

> ... current exchange rates already reflect *current expectations* about the future, while changes in the current exchange rate reflect primarily, *changes in these expectations* which, by definition, arise from new information.[1]

Frenkel's specification, equation 2, employs the difference between the *realized* interest differential and the *expected* differential; however, his argument implies that the news variable should be the change in the expected differential between $t-1$ and t. That is,

$$(2.2)\quad e_t = \phi\,(E_t\,(i - i^*)_{t,t+1} - E_{t-1}\,(i - i^*)_{t-1,t}).$$

IRP implies that the annualized one-month forward premium,

$$(2.3)\quad fp_t \equiv 12(f_t - s_t),$$

is equal to the interest differential expected to prevail during t through $t+1$,

$$(2.4)\quad fp_t = E_t\,(i - i^*)_{t,t+1},$$

where the term to maturity of the interest rates is equal to the holding period in fp. If this equality did not hold, riskless opportunities for profitable arbitrage would exist.[2] Thus, substituting the relevant forward premia from equation 2.4 for the expected interest differentials in equation 2.2 and then substituting this expression for the error term in equation 2.1, we obtain

$$(2.5)\quad s_t - f_{t-1} = \phi\,(\Delta fp_t),$$

which can be written in an estimable form as

$$(3)\quad s_t - f_{t-1} = \alpha + \beta\,\Delta fp_t + \omega_t.$$

[1]Frenkel (1981b), pp. 700–701, emphasis added. Frenkel notes (see footnote 31, p. 701) that Gustav Cassel, "the most recognized proponent of the purchasing power parity doctrine" also recognized this forward-looking aspect:

> The international valuation of the currency will, then generally show a tendency to anticipate events, so to speak, and become more an expression of the internal value that the currency is expected to possess in a few months, or perhaps in a year's time (Cassel 1930, pp. 149–50).

[2]This is known as the covered arbitrage condition. For example, if the $fp_t < (i_{US} - i^*_{UK})_t$, an investor could sell pounds and buy dollars at time t, use the proceeds to buy a U.S. security; by buying forward pounds at t, the investor removes any exchange rate risk and obtains a higher yield than he would have in U.K. securities. Since this yield differential is riskless, arbitrage should drive it to zero and, in the process, ensure the equality shown in equation 2.4. For a fuller discussion and many instructive examples, see Wood and Wood (1985), pp. 378ff.

Therefore, regression estimates of equations 2, 3 or 5 may be sensitive to changes in policy goals and regimes.

In particular, the hypotheses for real and inflation news summarized above are dependent on the monetary policy regime. For example, when the monetary authority targets monetary growth, interest rates will be determined by the private and public demand for loanable funds; unforeseen changes in that demand will cause changes in interest rates. Interest rates also will reflect private expectations about inflation. In such a monetary policy regime, the Fisher hypothesis holds, so that real interest rates are simply the difference between nominal interest rates and anticipated inflation; consequently, equation 4 holds, while equation 5 follows as an implication of equations 3 and 4.[9]

In contrast, consider a monetary policy regime of

[9]However, a critical caveat in evaluating equation 5 (or 5′, see below) is Fama's assertion that, when complete PPP does not hold, uncertainty and differential tastes combine "to strip the Fisher equation of its meaning" (1984, p. 323).

targeting interest rates.[10] Under such a policy stance, movements in interest rates are, to some extent, policy determined in the short run since changes in the nominal interest rate induce offsetting changes in the money supply through a policy-reaction feedback.

Consequently, changes in interest rates under a regime of targeting interest rates convey different information than do interest rate changes under a regime of targeting monetary aggregates. A real interest differential under interest-rate targeting cannot be closed by capital flows alone if the monetary authority chooses to maintain a particular nominal target rate which maintains the differential. Over time, an interest rate target below the market rate will increase the inflation differential. The adjustment process then depends totally upon the relative inflation rates to restore PPP. And, again, the risk premium embodied in the intercept should be smaller during an interest-rate regime due to the reduced short-run, interest-rate uncertainty.

This policy regime hypothesis can be tested by an F-test on the restriction implicit in both equation 3 and 5 that the coefficients — α, β, β_1, β_2 — are stable over changes in monetary policy regimes. The restriction is tested by adding intercept and slope dummy variables to get equations 3′ and 5′, computing the F-statistic on the change in the residuals between the estimates of the restricted and unrestricted equations:

$$(3') \quad s_t - f_{t-1} = \alpha_0 + \alpha_{01}D + \beta_0\Delta fp_t + \beta_{01} D\Delta fp_t + \omega_t'$$

$$(5') \quad s_t = f_{t-1} = \alpha_0 + \alpha_{01}D + \beta_1 \Delta(r_t - r_t^*) + \beta_{11}D\Delta(r_t - r_t^*) + \beta_2 \Delta(\pi_t - \pi_t^*) + \beta_{21}D\Delta(\pi_t - \pi_t^*) + \varepsilon',$$

$$\text{where } D = \begin{cases} 1 & \text{if October 1979} \leq t \leq \text{September 1982} \\ 0 & \text{otherwise.} \end{cases}$$

Summary of Testable Implications

The implications of the analysis in equations 3′ and 5′ are worth summarizing before reporting the estimation results. First, news about the real interest differential causes negative forecast errors, $s_t - f_{t-1}$, while changes in the inflation differential cause positive forecast errors. If there are periods dominated by relative volatility in inflation and other periods dominated by real yield volatility, then equation 3, which restricts the coefficients to equality, should be rejected by an F-test in comparison with equation 5 which does not restrict these coefficients to equality.

Second, the theory underlying equation 5 implies that news about the expected inflation differential will cause forecast errors, $s_t - f_{t-1}$, whose magnitude depends on the sensitivity of money demand to changes in the inflation rate. The coefficient should have the same sign as the change in the inflation differential. Given the shortness of the observation period — one month — the regression coefficient β_2 in equation 5 should be positive but may not be significant.

Third, since the interest rates (hence, forward premia) are assumed to be determined without a monetary policy reaction function in the analysis represented in equation 5, monetary policy based on interest-rate targets affects these hypotheses. If the monetary policy regime affects the market valuations, i.e., spot and forward exchange rates, hence forward-rate forecast errors, then the restrictions in equation 5 which are removed in equation 5′ will be rejected by an F-test on the improved fit of equation 5′ relative to equation 5.

Fourth, since it is well known that the variances of U.S. interest rates, both nominal and real, have been higher during monetary target regimes than alternative regimes, there is a greater likelihood of misforecasting interest rates under a monetary target regime.[11] The risk premium measured by the intercept, which primarily is determined by this risk, should be negative, larger and more significant during periods of monetary targeting than during periods of interest-rate targeting. This hypothesis can be tested by the significance of the intercept's dummy variable in equations 3′ or 5′.

Finally, under the efficient market hypothesis embodied in equations 3, 5, 3′ and 5′, the error terms should be serially uncorrelated. Correlation in the disturbance term implies incomplete use of past information and failure to exhaust profit opportunities. Alternatively, if markets are efficient, serially correlated residuals imply a misspecification of the estimating equation.

[10] Only two U.S. monetary policy regimes are distinguished in this study — the October 1979–September 1982 period and the remaining period before and after. Implicitly, this assumes that both the pre-October-1979 and the post-September-1982 periods are based on interest-rate targeting procedures; support for this characterization of these two periods is offered in Gilbert (1985), Kaufman (1982) and Rasche (1985). The foreign monetary policy stance might also be argued to be relevant; while this is a possibility for a refinement on the estimates reported in this study, there do not appear to have been substantial changes during the period 1974–83 in six of the eight countries. The policy procedures of six of the eight non-U.S. countries (excluding Italy and Netherlands) are reviewed in Johnson (1983).

[11] See Roley (1983) and Rasche (1985).

EMPIRICAL TESTS

The models specified in equations 1, 3, 5, 3′ and 5′ were estimated using monthly data from October 1973 through June 1985, using the U.S. dollar spot and one-month-forward prices of the currencies of Canada, France, Germany, Italy, Japan, the Netherlands, Switzerland and the United Kingdom. The tests are nested in that equation 3 is obtained from equation 1 by imposition of the efficient market hypothesis. Equation 1 also contains both the restriction to suppress the real interest rate vs. inflation rate decomposition and the restriction to suppress the effects of changing monetary policy regimes on the regression coefficients' values. We first test the simple efficient market hypothesis by estimating equation 1. Next, we estimate the simple news model with the change in the nominal forward premium, equation 3. This model contains both the nominal news and the policy regime restrictions above. We can then test these restrictions by estimating 5′, which is unrestricted and comparing it through F-tests with equations 5 and 3′. F-tests on equation 5′ vs. equation 5 and 5′ vs. 3′ determine, respectively, whether the policy regime or nominal forward premium restrictions can be rejected.

Data

The spot and 30-day forward exchange rates used in the estimates are New York opening market (10 a.m. midpoints) for the last business day of the month as compiled by the Bank of America. The change in the real interest differential was obtained from the change in the forward premium: First, the forward premium was converted to an annualized rate; the change in this annualized forward premium is the news — that is, the change in the expected nominal interest differential. Second, an expected annualized inflation rate for the one-month horizon was computed for each country from its monthly CPI series.[12] The change in the differential, U.S. minus foreign inflation, is the change in the inflation differential used in estimating equations 5 and 5′. The change in the real interest differential is then the change in the annualized, nominal, one-month-forward premium minus the change in the expected inflation differential.

[12]Clemens Kool of Erasmus University computed this series using a multi-state Kalman filter. A simple Kalman filter is a forecasting method based on assumptions about the forecasted variable's relation to current and lagged data on itself and or other series. A multi-state Kalman filter allows this relation to vary according to a feedback or adaptive error loop; the multi-state modifier refers to the alternative sets of assumed weights. A concise description and illustrative example are contained in the statistical appendix to Bomhoff and Korteweg (1983).

Tests of Forward Market Efficiency

Table 1 reports the results of estimating equation 1 during the full sample period, October 1973 through June 1985. For six of the eight currencies considered, market efficiency is not rejected; for Japan and Switzerland, however, the market efficiency hypothesis is rejected at the 5 percent level. For all eight, the Durbin-Watson statistic indicates that hypothesis of serially uncorrelated disturbances is not rejected. Thus, except for Japan and Switzerland, the results in table 1 indicate that the news model specified in equation 3 is an appropriate empirical model.

For Japan and Switzerland, equation 1 was reestimated by subperiods before, during and since the U.S. monetary aggregate target regime of October 1979 through September 1982. For each country, the hypothesis of serially uncorrelated residuals was not rejected in any subperiod. For each of the subperiods, the efficient market hypotheses bearing on the coefficients for Switzerland were not rejected. For Japan, the earlier two subperiod estimates do not reject market efficiency, but the recent subperiod rejects market efficiency both in terms of a significant intercept and the deviation from unity of the lagged forward rate coefficient.[13]

Consequently, for neither Switzerland nor Japan is the estimation of equation 3 justified since equation 3 is derived from equation 1 assuming a unit coefficient on f_{t-1}. Yet, equation 3′ or equation 5′ may be justified for Switzerland since the dummy variables can account for the nonstable coefficient. For Japan, the failure of the efficient market hypothesis in the last subperiod is not offset by any of our variables, and it is consistent with this failure that Japan rejects each of the specifications equations 3′, 5 and 5′ as reported in tables 2 and 3.

Tests of News Model with U.S. Monetary Regimes Not Distinguished

Table 2 reports the results of estimating equation 3, the news model with the change in the nominal forward premium, over the full period, October 1973–June 1985. In sharp contrast to the results in table 1, which support this specification, the estimates uniformly reject this model: no coefficient is significant at

[13]The October 1982–June 1985 estimates for Japan are very curious. The estimated intercept is huge in comparison with the earlier-period Japanese estimates, the Swiss estimates or any of the estimates in table 1:

$$\alpha = -1.192 \text{ (s.e. } = 0.548), \beta = 0.783 \text{ (s.e. } = 0.100).$$

Table 1

Tests of Forward Exchange Market Efficiency for U.S. Dollar, October 1973–June 1985 (U.S. Monetary Regimes Not Distinguished)

	Coefficients[1]		Summary Statistics		Test
Currency	Intercept	f_{t-1}	$\bar{R}^2$	DW	F^2
Canada	−0.002	0.998	0.981	2.16	0.730
	(0.002)	(0.012)			
France	−0.002	1.001	0.985	2.07	0.573
	(0.018)	(0.010)			
Germany	−0.020	0.981	0.954	2.03	1.800
	(0.016)	(0.018)			
Italy	0.012	1.002	0.992	1.87	0.500
	(0.052)	(0.007)			
Japan	−0.298	0.946	0.940	1.80	3.734*
	(0.112)*	(0.020)*			
Netherlands	−0.013	0.991	0.957	2.01	1.486
	(0.017)	(0.018)			
Switzerland	−0.034	0.961	0.962	1.92	3.537*
	(0.013)*	(0.016)*			
United Kingdom	0.001	0.994	0.974	1.82	0.483
	(0.009)	(0.014)			

[1]Standard errors of estimated coefficients appear in parentheses; asterisks indicate rejection at 5 percent level of individual efficient market hypotheses — intercept is zero, slope coefficient = 1.0.

[2]F-test of joint efficient market hypothesis that intercept is zero *and* slope coefficient is unity; asterisk indicates rejection at 5 percent level.

any reasonable confidence level and the adjusted R^2 is negative for six of the eight currencies tested. Consistent with the efficient market hypothesis, however, the hypothesis of serially uncorrelated disturbances is not rejected. Nonetheless, the results require an investigation of alternative explanations for this model's uniform failure.

Decomposition of Nominal Forward Premium

Also reported in table 2 is the F-statistic for testing whether decomposing the change in the nominal forward premium into innovations in its expected real and inflation components is statistically warranted. The F-statistic is obtained from the difference in the explanatory power of equation 5 with respect to equation 3; the critical value for rejecting the restriction in equation 3 (that β_1, β_2 in equation 5 are equal) is 3.92. Only the Netherlands result rejects the restriction.

Tests of News Model with U.S. Monetary Regimes Distinguished

As discussed above, the U.S. monetary policy regime can be expected to affect the relationship between the dollar's exchange rates and U.S.-foreign interest differentials. Thus, the statistical results reported in table 2 may be invalid because they do not distinguish changes in the U.S. monetary policy stance. To test for such policy regime effects, equations 3′ and 5′, were estimated to isolate the period of U.S. monetary aggregate targeting, from October 1979 to September 1982, with slope and intercept dummies.

Table 3 reports estimates of equation 5′ and the F-statistics to test the effect of monetary regime changes and the equality restriction implicit in equation 3′ and removed in equation 5′. The estimates present a substantial contrast to those in table 2. Canada and Italy reject the nominal forward premium restriction (last

Table 2

Tests of News Model Using Change in Nominal Annualized Forward Premium on U.S. Dollar, October 1973–June 1985 (U.S. Monetary Regimes Not Distinguished)

	Coefficients[1]		Summary Statistics			Test
Currency	Intercept	Δfp	$\bar{R}^2$	DW	F	F[2]
Canada	−0.001 (0.001)	−0.069 (0.102)	−0.004	2.16	0.455	2.059
France	−0.003 (0.003)	−0.076 (0.066)	0.002	2.04	1.347	2.937
Germany	−0.004 (0.003)	−0.146 (0.188)	−0.003	2.06	0.600	1.898
Italy	−0.002 (0.002)	−0.016 (0.027)	−0.005	1.86	0.332	0.053
Japan	−0.002 (0.003)	−0.017 (0.035)	−0.006	1.80	0.236	1.254
Netherlands	−0.005 (0.003)	0.031 (0.031)	0.000	2.03	1.004	4.164*
Switzerland	−0.004 (0.003)	−0.011 (0.179)	−0.007	1.92	0.004	0.326
United Kingdom	−0.002 (0.003)	0.006 (0.123)	−0.007	1.83	0.002	1.888

[1]Standard errors of estimated coefficients appear in parentheses.

[2]F-statistic for testing the equality restriction on the coefficients of the change in the real and the inflation differentials (components of the change in the nominal forward premium); asterisk indicates rejection at 5 percent level.

column, F-test) but, in contrast to table 2, the Netherlands does not when the U.S. monetary regime shift is accounted for. Considering the appropriate specification, equations 3′ or 5′, six of the eight equations are significant in terms of their overall fit (F-statistics) at the 5 percent level, France is significant at the 6 percent level, and seven of eight countries reject the restriction of stable coefficients across monetary regime changes at the 10 percent level or better. Only Japan fails the F-test for the significance of the model.

In terms of the individual coefficients, six of the eight countries evidence a significant negative risk premium (10 percent or better) during the U.S. monetary aggregate targeting period, while the intercept is uniformly nonsignificant during the other U.S. monetary policy regime, October 1973–September 1979 and October 1982–June 1985. The impact of the different regimes is also notable in the slope interaction dummy. The coefficient on the change in the real forward premium is negative and significant for Canada, Germany, the Netherlands, Switzerland and the United Kingdom. For Germany, Switzerland and the United Kingdom, this entails a switch from a positive and significant coefficient during the U.S. non-monetary targeting regime.

Thus, for each of the seven currencies for which the market efficiency criteria are met, the U.S. monetary policy regime has a significant effect on the errors in the forward rate forecasts. More specifically, two generalizations can be advanced based on the results in table 3. First, the greater interest rate volatility during U.S. monetary aggregate targeting shows up in a significant risk premium tending to strengthen the dollar against six of the eight currencies. Second, given the failure to reject the nominal forward premium restriction of equation 3′, the negative significance of the slope dummy implies that the interest differential news was primarily interpreted as an increase in the inflation differential during U.S. non-monetary aggregate targeting periods and as an increase in real interest differentials during U.S. monetary aggregate targeting. In other words, the dollar appreciated along with unanticipated increases in the forward premium during October 1979 to September 1982, but depreciated

Table 3

Tests of News Model Using Unrestricted Specification, October 1973–June 1985 (U.S. Monetary Regimes Distinguished)

Currency	Coefficients[1] Intercept	DI[2]	$\Delta(r-r^*)$	Dr[2]	$\Delta(\pi-\pi^*)$	Dπ[2]	Summary Statistics $\bar{R}^2$	DW	F	Tests F[3]	F[4]
Canada	−0.002	0.000	0.299	−0.436	−0.343	−0.776	0.056	2.19	2.728*	3.658*	4.027*
	(0.001)	(0.003)	(0.188)	(0.223)+	(0.209)+	(0.253)*					
France	0.001	−0.015	−0.023	−0.077	−0.368	0.211	0.045	2.12	2.297*	2.352+	2.133
	(0.003)	(0.006)*	(0.105)	(0.135)	(0.193)+	(0.282)					
Germany	−0.001	−0.014	0.540	−1.137	−0.311	−0.980	0.081	2.10	3.445**	4.837**	0.605
	(0.003)	(0.006)*	(0.299)*	(0.382)*	(0.395)	(0.511)+					
Italy	0.000	−0.011	0.012	−0.067	0.139	−0.568	0.059	1.92	2.731*	4.415**	3.990*
	(0.003)	(0.006)+	(0.033)	(0.056)	(0.100)	(0.181)*					
Japan	0.001	−0.012	0.029	−0.255	0.163	−0.434	0.016	1.89	1.443	1.899	1.093
	(0.003)	(0.006)+	(0.044)	(0.200)	(0.124)	(0.292)					
Netherlands	−0.001	−0.013	0.047	−0.816	−0.280	−0.536	0.107	2.06	4.324**	5.313**	2.098
	(0.003)	(0.006)*	(0.029)	(0.243)*	(0.161)+	(0.343)					
Switzerland	0.001	−0.018	0.433	−1.217	0.506	−1.191	0.088	2.03	3.688**	6.025**	0.162
	(0.004)	(0.007)*	(0.218)*	(0.356)*	(0.263)+	(0.437)*					
United Kingdom	−0.001	−0.006	0.382	−0.950	0.319	−1.087	0.097	1.79	3.979**	5.933**	1.052
	(0.003)	(0.006)	(0.149)*	(0.238)*	(0.180)+	(0.296)*					

[1]Standard errors of estimated coefficients appear in parentheses; asterisk indicates significance at 5 percent level and plus sign indicates significance at 10 percent level.

[2]DI, Dr and Dπ equal 1.0 during period of U.S. monetary-target policy regime, October 1979–September 1982 and zero otherwise.

[3]F-statistic for testing restriction that coefficients are stable across different monetary regimes; double asterisk indicates rejection at 1 percent level, asterisk indicates rejection at 5 percent level, and plus indicates rejection at 10 percent level.

[4]F-statistic for testing the equality restriction on the coefficients of the change in the real and the inflation differentials (components of the change in the nominal forward premium); asterisk indicates rejection at 5 percent level, plus indicates rejection at 10 percent level.

with such news during the rest of the floating rate period. This is consistent with Frenkel's (1981a) results for 1973–79. Finally, the Durbin-Watson statistics in table 3 do not indicate serial correlation in the residuals, consistent with the maintained hypothesis of market efficiency.

There remain two puzzling results: (1) The estimated coefficients of the change in the inflation differential during the monetary regime are generally negative, refuting the hypothesis embodied in equation 5; this negative coefficient is significant at the 10 percent level or better in five countries. (2) Moreover, the decomposition of the nominal interest differential is significant only for Canada and Italy. This irrelevance of the distinction between real and nominal interest differentials may simply be a confirmation of Fama's (1984) assertion that, with risk aversion or without PPP, the Fisher equation does not hold (see footnote 9). Indeed, for six of the eight currencies, the F-test does not reject the implicit restriction of equality of changes in the nominal interest differential's two components displayed in table 3.

The Implications of Monetary Regimes: A Closer Look

The negative coefficient on the inflation differential during the 1979–82 monetary regime is both pervasive and puzzling. Two possible explanations are worth considering. First, the one-month horizon of the estimated, anticipated CPI inflation rates used in estimating equation 5′ may be too short, or the estimated expected inflation series simply may be bad proxies.

Second, the market may have determined that the U.S. monetary authority and the administration were committed to lowering the U.S. inflation rate. Consequently, a short-term increase in the U.S. expected inflation rate would lead market participants to expect a tightening of monetary growth.[14] If so, a short-term increase in U.S. inflation would lead to increases in the U.S. real interest rate as the market anticipated the monetary authority's reaction. This explanation, consistent with research by Cornell (1982), has not been tested here, but it is consistent with the decomposition of changes in the nominal interest differential generally not increasing the explanatory power of the equation for six of the eight currencies.[15]

CONCLUSION

We have tested the efficiency of forward exchange markets for the dollar against eight major currencies during the floating period. The regression estimates clearly demonstrate that failing to account for changes in the policy procedures of the U.S. monetary authority entails misspecification. Monetary regime changes alter the risk premia that market participants require on forward contracts and affect the direction of errors implied by nominal and real news, that is, unforeseen events occurring between the time of contract and its maturity. The implications of the standard model of exchange rate behavior were substantiated for nominal news under a monetary target regime, but its implication for inflation differentials was refuted. While a closer modeling of the policy procedure may explain this rejection, it remains a prominent puzzle in this study. Nonetheless, one interpretation of these results is that market participants regarded the U.S. monetary policy regime of 1979–82 as anti-inflationary. If this is correct, it follows that credible goals of monetary policy may be as significant for market participants as the mechanical details of that policy's execution.

[14]The U.S. CPI inflation rate was 13.3 percent in 1979, 12.4 percent in 1980, 8.9 percent in 1981 and 3.9 percent in 1982. There is also some support for this view in the impact of lagged reserve accounting during the monetary targeting period. As Kaufman (1982) notes, this results in more volatility of both money *and* interest rates since a decision to maintain a target growth path when the money supply has exceeded the path requires a subsequent reduction of reserve growth. Since banks already will have increased their required reserves, real rates will vary with the money supply errors and, perhaps, short-run inflation expectations.

[15]Cornell (1982) finds that unexpected monetary supply increases are correlated with an appreciation in the dollar, not the depreciation that an anticipated simple link with increased inflation would imply. Cornell suggests that the explanation is an anticipated policy reaction, a tightening of the money supply growth rate.

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How Federal Farm Spending Distorts Measures of Economic Activity

John A. Tatom

DURING the 1980s, federal purchases of farm products by the Commodity Credit Corporation (CCC) have exhibited relatively large quarterly swings that have significantly affected how we interpret economic developments.[1] Although these purchases increase the government's inventory of farm products, they are treated as final sales to the government, instead of inventory transactions, in the National Income and Product Accounts (NIPA). As a result, a CCC purchase increases federal purchases and final sales in the economy and reduces measured investment in farm inventory. Similar private sector transactions, which redistribute farm products from one owner to another, result in offsetting changes in farm and business inventory; these transactions affect neither business inventory investment nor final sales.

This article explains the impact of CCC purchases and examines the distortions that they can produce in quarter-to-quarter movements of some important NIPA measures. It shows that adjusting for the effect of CCC purchases can alter conclusions about the short-term performance and outlook for federal purchases, the farm sector and aggregate production and employment. The largest swings in CCC purchases on record were recorded at the end of 1985 and early this year; hence, these recent swings have had the greatest impact on measures of inventory investment, federal purchases and overall final sales. A more useful perspective on NIPA measures can be obtained by adjusting these measures during quarters when large changes in CCC purchases occur.

CCC PURCHASES, SALES AND INVENTORY CHANGES

The Commodity Credit Corporation, established in 1933 as part of the Department of Agriculture, carries out the federal government's price support programs.[2] These programs include both "nonrecourse loans" and direct purchases of farm products. The former are called nonrecourse loans because the farmer is free to deliver the pledged crop, which serves as collateral, in order to settle the loan.[3] The price of the commodity at which the loan is advanced is called the loan rate; it establishes a minimum price for the commodity. When the government makes such a loan, the transaction is treated in the NIPA as a purchase of farm products. As a result, these loans increase federal purchases and reduce farm inventory holdings. Repayment of the loan reverses these accounting entries.[4]

Direct purchases of farm products are treated in the

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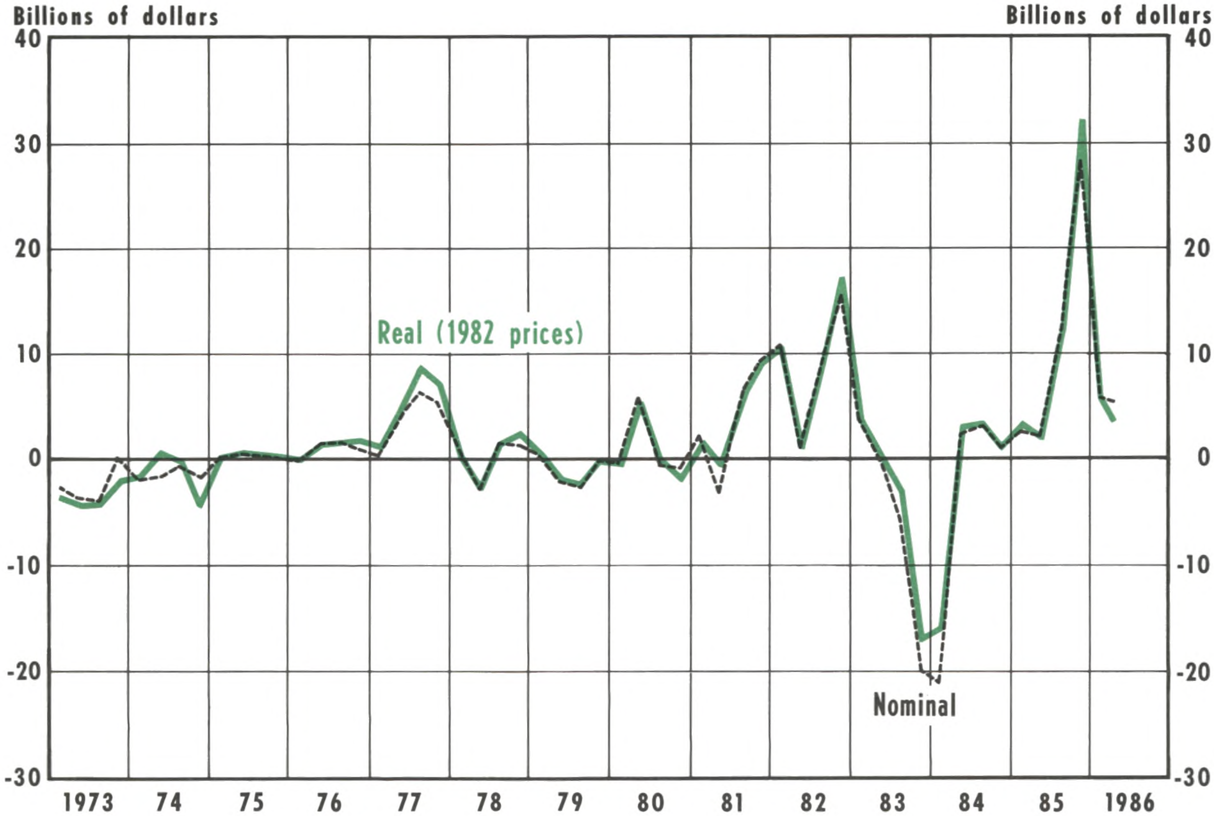
[1]The significance of such swings, especially as a major source of changes in federal purchases, was first noted by the Bureau of Economic Analysis (1982).

[2]More extensive discussion of the CCC can be found in the Council of Economic Advisers (1986), Herman (1978), Bureau of Economic Analysis (1982) and Wakefield (1986). The former also details other features of U.S. agricultural policy.

[3]Nonrecourse loans to farmers are based on the government-set loan rate for each farm product and the amount of the current or past product pledged against the loan as collateral. If the producer-borrower cannot sell his product for more than the loan rate plus the accumulated storage costs and interest on the loan, the farmer forfeits the pledged crop and the loan obligation is discharged. The farm products that are covered by the loan program include wheat, corn, barley, oats, rice, cotton, honey, peanuts, sorghum, soybeans, rye, tobacco and sugar.

[4]Even when the farmer pays off the loan, he reaps a benefit in the form of a short-term credit subsidy, since the interest rate on such loans is less than market rates. The CCC also supports prices of farm products by directly purchasing certain products at official support prices when such prices exceed market levels. Chief among these are such dairy products as cheese, butter and dry milk.

Chart 1
CCC Purchases



exact same way in the NIPA. Thus, commodity loans and direct commodity purchases by the federal government result in offsetting changes in federal purchases of goods and services and business (farm) inventory investment. GNP is unaffected by the transactions because they result in no change in production.[5]

Chart 1 shows both nominal and real (1982 prices) CCC inventory purchases from 1973 to the second quarter of 1986. Although the nominal purchases appear small relative to current GNP of over $4 trillion, the quarter-to-quarter swings are sometimes quite large in comparison to GNP movements. For example, in the fourth quarter of 1985, such purchases rose $20.8 billion, or 36.5 percent of the total increase in GNP during the same quarter. It is also evident from the chart that movements in CCC purchases have become substantially larger in the 1980s, with the biggest swings occurring at the end of 1985 and in early 1986. In part, these increased fluctuations reflect the growing role of federal farm programs.

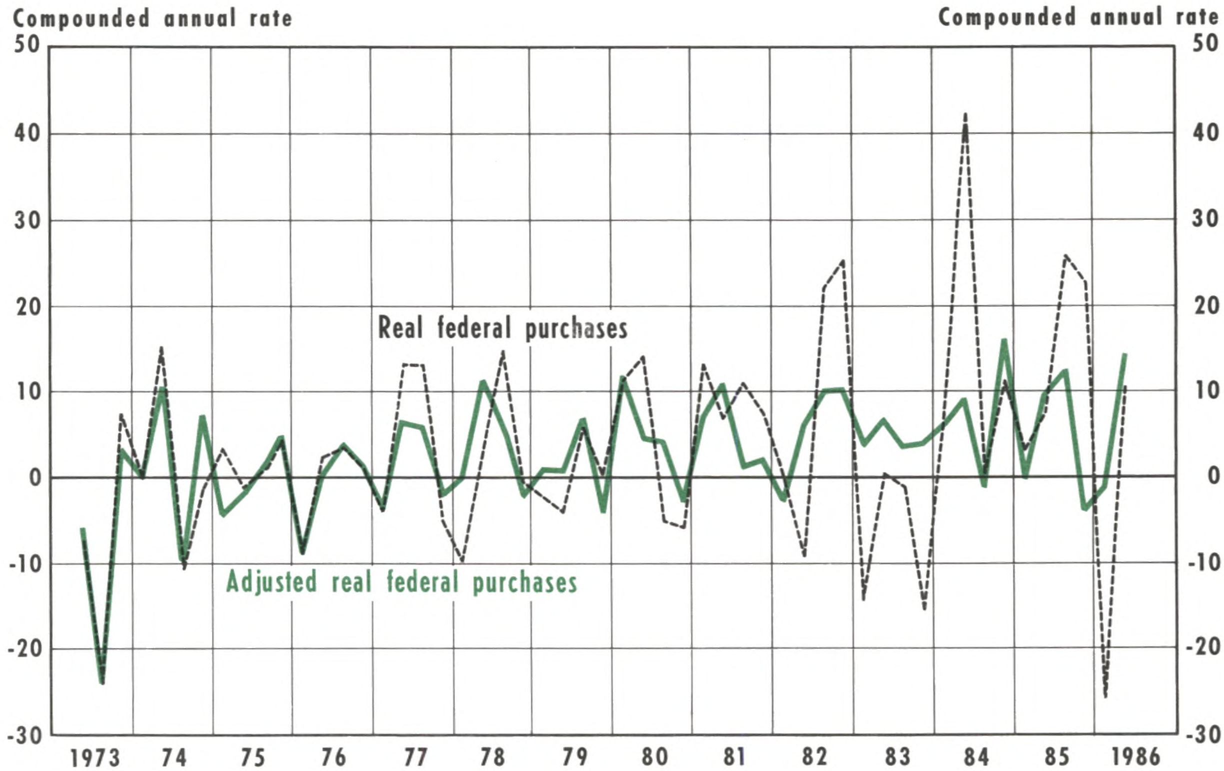
CCC AND FEDERAL PURCHASES OF GOODS AND SERVICES

Quarterly movements in CCC purchases have had a sizable impact on the pattern of growth of federal purchases during some quarters in the 1980s. Chart 2 shows the growth rates of real federal purchases and adjusted real federal purchases (which exclude CCC

[5]The independence of GNP from CCC purchases is based on two assumptions: (1) that the coverage, timing and seasonal adjustment of changes in farm inventory and CCC purchases are consistent and (2) that farmers, in general, cannot or do not respond to CCC purchases within the quarter by altering production. The former point has been made by the Bureau of Economic Analysis (1982). These second-order considerations are ignored below in order to focus solely on the measurement principles involved.

Chart 2

The Growth Rate of Real Federal Purchases with and without CCC Purchases



purchases) since 1973.[6] In the 1980s, the difference in the growth rates often has been quite large and more variable. Since 1980, the federal government generally has been accumulating inventory of farm products, but in 1983 and early 1984, the Payment-In-Kind (PIK) program led to large sales for four quarters.[7] These swings in CCC purchases had a major impact on the growth rate of federal purchases, generally depressing it in 1983 and early 1984 and subsequently raising it. These swings make it difficult for analysts to interpret trends in federal spending.

Another coincidental effect of CCC purchases in recent years has been to raise the growth rate of federal purchases during recession periods, while depressing the growth of federal purchases during the initial stages of expansions. This effect has resulted in the appearance of a negative relationship between GNP and federal purchases, a relationship that disappears when federal purchases are adjusted for CCC purchases. For example, from I/1980 to II/1986, the correlation between the growth rate of real federal purchases of goods and services including CCC purchases and of real GNP is negative (−0.15); when real CCC purchases are omitted from government purchases, however, the correlation is positive (0.04). While neither correlation is statistically significant, distortions caused by volatile CCC purchases can bias statistical tests of fiscal policy's general effectiveness.

CCC PURCHASES AND CHANGES IN FARM INVENTORY

Federal purchases of farm products are offset in the

[6]Since nominal and real CCC inventory changes are not substantially different over the period since 1973, attention throughout this article is focused on real measures. Movements in the nominal counterparts of real measures provide no additional insight and so are ignored here.

[7]A description and analysis of the PIK program that was in effect in 1983 and early 1984 can be found in Belongia (1983) and Rosine (1984).

Table 1

The Change in Farm Inventory and CCC Purchases (billions of dollars, 1982 prices)

	CCC purchases	Change in farm inventory	Annual mean/ standard deviation	Change in farm inventory and CCC	Annual mean/ standard deviation
I/1980	$-0.3	$-5.0		$-5.3	
II	5.5	-7.0	$-4.7	-1.5	$-3.9
III	-0.2	-10.5	6.09	-10.7	5.37
IV	-2.0	3.8		1.8	
I/1981	1.6	4.6		6.2	
II	-0.8	11.2	4.9	10.4	8.7
III	5.5	5.0	5.11	10.5	2.09
IV	9.1	-1.3		7.8	
I/1982	10.8	-4.1		6.7	
II	0.7	4.0	-1.5	4.7	7.7
III	7.9	3.2	6.16	11.1	2.71
IV	17.2	-8.9		8.3	
I/1983	3.8	-9.1		-5.3	
II	-0.1	-6.9	-6.3	-7.0	-10.5
III	-3.1	-15.7	9.32	-18.8	6.01
IV	-17.2	6.5		-10.7	
I/1984	-15.9	16.4		0.5	
II	3.1	1.8	4.9	4.9	2.7
III	3.4	1.3	7.72	4.7	2.4
IV	0.8	0.0		0.8	
I/1985	3.2	6.4		9.6	
II	2.0	7.8	-2.0	9.8	10.3
III	11.5	-0.7	13.43	10.8	0.7
IV	32.3	-21.3		11.0	
I/1986	6.4	2.9		9.3	
II	4.5	4.1	—	8.6	—

GNP accounts by reductions in farm inventory.[8] Thus, CCC purchases can distort the short-run interpretation of changes in farm and business inventory. When the CCC purchases (sells) farm goods, farm and business inventory investment falls (rises), giving the appearance of an inventory change. Of course, such an appearance is deceptive; in fact, inventory holdings have simply moved from private to federal government ownership, or vice versa.

Table 1 shows quarterly real CCC purchases and changes in both real farm inventory and real farm inventory plus real CCC purchases since 1979.[9] The mean and standard deviation of each series also are shown for each year. The pattern of changes in the overall measure of farm inventory is much smoother when CCC purchases are included than when they are not. This is especially true when relatively large changes in CCC purchases occur. At these times, farm inventory investment swings widely in the opposite direction, such as in IV/1982, IV/1983, I/1984 and the

[8]An inverse relationship between business inventory investment and government purchases of goods has been noted by Weidenbaum (1959) and (1961). His analysis emphasizes the time pattern of production and delivery and the NIPA accounting of such programs. The implied lack of a contemporaneous relationship of GNP and such spending was first pointed out in these articles.

[9]For the period shown in table 1, the correlation between changes in CCC purchases (1982 prices) and changes in farm inventory investment is −0.56, which is statistically significant at the 1 percent level.

end of 1985. The standard deviation for farm inventory investment each year is sharply higher than that for the total farm product inventory change. This occurs because the movements of CCC purchases are offset by opposite movements in farm inventory purchases. Of course, this smoothing effect also occurs for the overall change in inventory — the sum of business (non-farm and farm) inventory change and CCC purchases.

CCC PURCHASES AND FINAL SALES

While federal purchases of farm products do not affect GNP — the value of final goods and services produced in the economy — they do affect the measurement of final sales, which equals GNP less the change in business inventory.[10] Analysts often focus on final sales in order to assess the strength and outlook for income, output and employment. Assessments of final sales are important both because inventory and production decisions are based on expectations of such sales and because unexpected changes in sales are absorbed by inventory fluctuations. Thus, movements in final sales relative to production provide information on future production changes and can give rise to an inventory cycle.[11] When sales are less than production, for example, the unsold products increase inventory. If the rise in inventory is undesired and unplanned, it will be eliminated by reducing production growth temporarily relative to that of expected sales. Moreover, if movements in GNP reflect temporary changes in production to adjust inventory, final sales can be a more useful gauge of the outlook than current production or GNP.

CCC purchases have substantial quarter-to-quarter effects on the measurement of final sales. This occurs because such purchases affect the change in business inventory but leave GNP unaffected. When CCC purchases increase, for example, measured final sales tend to rise because business (farm) inventory declines. Yet such purchases simply represent another way of holding farm inventory, not a significant increase in overall spending on goods and services that will likely lead to increased production. Thus, if the change in business inventory is adjusted to include CCC purchases, the adjusted final sales measure obtained can more closely gauge the actual final purchases of goods and services by consumers, business, government and foreign purchasers. Chart 3 shows real final sales growth both without an adjustment and with CCC purchases subtracted from final sales.

The largest differences in the growth of final sales, adjusted for CCC purchases, occur after 1981. In the second half of 1982, relatively large CCC purchases contributed to final sales growth. From the second to the fourth quarter of 1982, real final sales expanded at a 2.1 percent rate, higher than the 1.1 percent rate for adjusted real final sales. Subsequent reductions in the government's holding of farm product inventory through the PIK program led to an understatement of final sales growth. From the fourth quarter of 1982 to the fourth quarter of 1983, real final sales expanded at a 3.7 percent rate, but this was below the 4.8 percent rate of adjusted real final sales growth. In effect, the transfer of farm product inventory from the government to the private sector appeared only as a net business inventory change, which understated the growth of final sales. Of course, these periods match the end of the 1981–82 recession and early part of the current expansion. Thus, the cyclical swing in measured final sales growth understates the actual acceleration in adjusted final sales that took place.

The most recent CCC purchases, especially in the fourth quarter of 1985, are the largest on record. In the second quarter of 1985 and the second quarter of 1986, real CCC purchases were $2 billion and $4.5 billion, respectively. Thus, in each quarter, the final sales measure was little affected by CCC purchases; over the whole year, real final sales and real final sales adjusted for CCC purchases rose 2.7 and 2.6 percent, respectively. Moreover, the pace of overall inventory investment was about the same in each quarter, so that real GNP grew at about the same rate over the year.

But the patterns of real GNP, real final sales and adjusted real final sales were quite different during the year. Table 2 shows these growth rates. Both final sales series show that production grew faster than sales in

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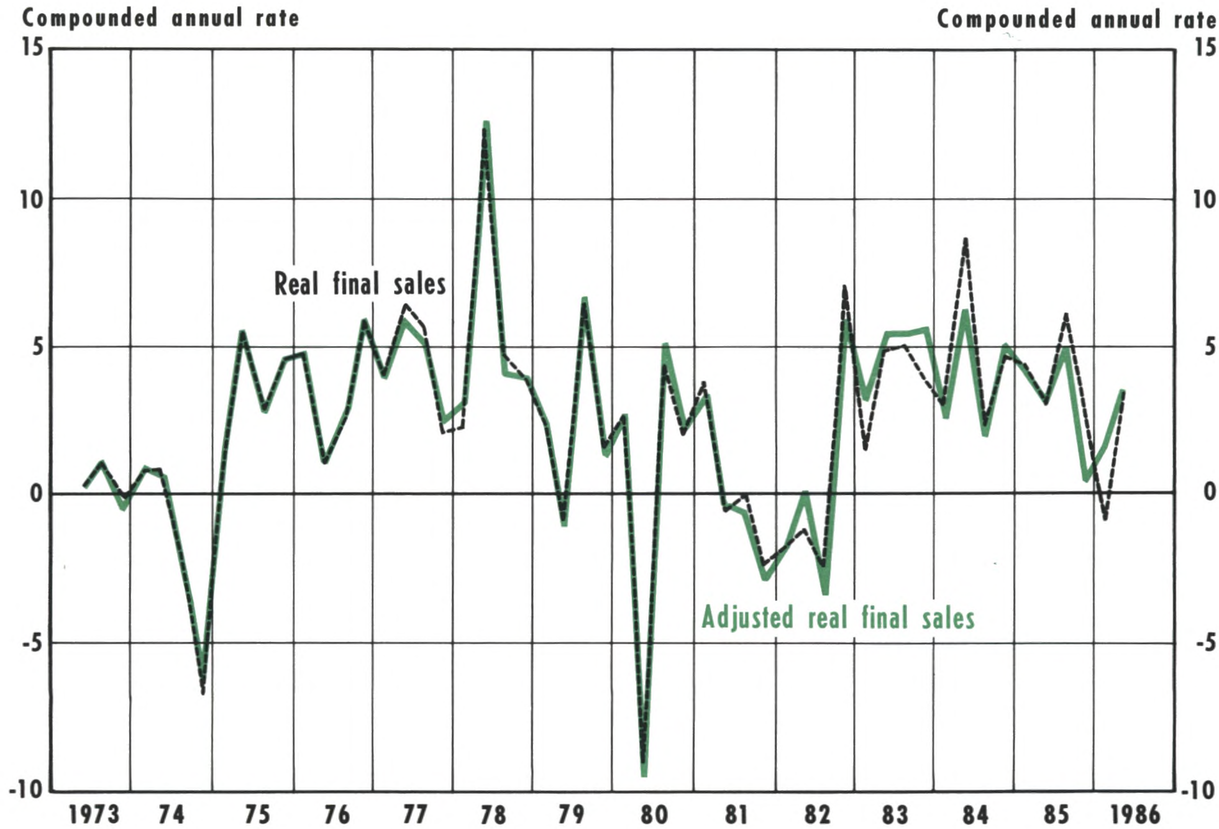
[10] While the assumed independence of CCC purchases and farm output *within the quarter* seems satisfactory, it might be argued that such purchases contribute to higher farm output than would otherwise occur. To test these views, "Granger-causality" tests were conducted on the quarterly change in farm sector output and the change in CCC purchases, both in 1982 prices, for the period II/1973 to II/1986. Optimal lags on the lagged dependent variable were chosen via sequential F-tests. The results indicate "bidirectional causality": past CCC purchases negatively and significantly affect farm output; past changes in farm output positively and significantly raise CCC purchases. When the contemporaneous value of the change in CCC purchases is included in the farm output equation, there is no significant past CCC effect and the contemporaneous CCC term is not significant for lags on the change in CCC purchases up to 10 quarters earlier.

[11] The inventory cycle and its significance in U.S. business cycles from 1948 to 1976 is discussed in Tatom (1977).

Chart 3

CCC Purchases and Real Final Sales Growth



the last quarter of 1985 and first quarter of 1986. So, not surprisingly, production growth slowed temporarily in the second quarter of 1986 to eliminate excess inventory. Both final sales series also show that sales growth accelerated in the second quarter of 1986.

The principal differences in table 2 are that sales growth in 1986 was stronger according to the adjusted series and that it accelerated for two quarters rather than one. The stronger sales growth on an adjusted basis suggests stronger growth in aggregate demand and more incentive for firms to increase production and employment than the unadjusted data indicate. Also, the second quarter acceleration in final sales appears less likely to be a fluke using the adjusted series. The acceleration simply reinforces the pattern set in the previous quarter, instead of appearing to be the first sign of positive sales growth since the end of 1985, as indicated in the unadjusted data.

SUMMARY

While movements in CCC purchases can be relatively large, they have had no major effects on final sales and other NIPA measures until the past few years. During recent years, the pattern of CCC purchases has had relatively large effects on measured inventory change, federal purchases and expenditures, and final sales. In 1982 and 1983, the effect was to raise the growth of both federal spending and final sales during the last two quarters of the recession and to lower their growth in the first five quarters of the subsequent expansion. More recently, record net purchases by the CCC in the last half of 1985 have given rise to a distorted pattern of sales growth, suggesting generally weaker sales than the adjusted data indicate. Analysts who focus on unadjusted data, accordingly, would understate the recent strength of aggregate demand and the short-run prospects for growth.

Table 2
Growth Rates of GNP and Final Sales over the Previous Year

Quarter ending	Real GNP	Real final sales	Final sales less CCC purchases
III/1985	4.1%	6.1%	5.0%
IV/1985	2.1	2.7	0.4
I/1986	3.8	−1.3	1.6
II/1986	0.6	3.4	3.6
II/1985–II/1986	2.6	2.7	2.6

For policy purposes, fluctuations in CCC purchases can distort quarter-to-quarter movements in important NIPA measures, providing a misleading indication of the strength or weakness of federal spending, farm inventory investment and final sales. Faced with such distortions, analysts will find it useful to take more care in accounting for these quarterly movements in CCC purchases and their effects on key measures of economic performance.

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