Permanent Income, Current Income, and Consumption

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ABSTRACT

This paper reexamines the consistency of the permanent income hypothesis with aggregate, post-war, United States data. The permanent income hypothesis is nested within a more general model in which a fraction of income accrues to individuals who consume their current income rather than their permanent income. This fraction is estimated to be 40 or 50 percent, indicating a substantial departure from the permanent income hypothesis. This finding is robust to various statistical problems that have plagued previous work, such as time aggregation, and cannot be easily explained by appealing to changes in the real interest rate or to non-separabilities in the utility function.

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1. Introduction

During the past decade, much effort has been directed at the question of whether the response of consumption to income is consistent with the permanent income hypothesis. Hall (1978) showed that a central implication of the theory is that consumption should follow a random walk. He argued that, to a first approximation, postwar U.S. data are consistent with this implication. In contrast, Flavin (1981) reported that consumption is "excessively sensitive" to income, a conclusion that has been widely interpreted as evidence that liquidity constraints are important for understanding consumer spending (Dornbusch and Fischer, 1987). Yet Mankiw and Shapiro (1985) showed that Flavin's procedure for testing the permanent income hypothesis is severely biased toward rejection if income has approximately a unit root. Nelson (1987) has recently reappraised the evidence on the permanent income hypothesis, and concludes that it is generally favorable.

Other recent research has examined the permanent income theory from a different point of view. Campbell (1987) studies the implications of the theory for savings behavior, while Campbell and Deaton (1987) and West (1986), following Deaton (1986), look at its implications for the "smoothness" (the standard deviation of the change) of consumption. These papers argue that while some of the qualitative implications of the model are fulfilled, consumption appears to be too smooth and there is weak evidence that saving moves too little to be consistent with the theory.

The first goal of this paper is to provide a simple framework for understanding these disparate results. We nest the permanent income

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hypothesis in a more general model in which some fraction of income λ accrues to individuals who consume their current income, while the remainder $(1-\lambda)$ accrues to individuals who consume their permanent income[1]. We show that a value of λ greater than zero can generate excess sensitivity of consumption to income in the sense of Flavin and insufficient variability of saving as found by Campbell. It may also imply excess smoothness of consumption if the stochastic process generating income is highly persistent.

We show how to estimate λ and test the permanent income hypothesis that $\lambda = 0$, using an instrumental variables approach. Our test is valid whether or not income has a unit root, and it is more powerful than the standard unrestricted test for consumption following a random walk. By lagging our instruments two periods, we are able to avoid econometric difficulties which would otherwise be created by time aggregation of our data. We also show how to test our framework against an even more general time-series representation for consumption and income, for example a disequilibrium "error-correction" model of the type proposed by Davidson, Hendry, Srba and Yeo (1978) and Davidson and Hendry (1981).

The second goal of this paper is to generalize the above approach to handle alternative versions of the permanent income hypothesis. We can allow for changes in the real interest rate, as in Grossman and Shiller (1981), Mankiw (1981), Hansen and Singleton (1983), Bean (1986) and Hall (1987). We can also allow for non-separability in the utility function between consumption and other goods. Following previous work, we examine

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[1] This model has also been studied by Hall and Mishkin (1981), Hayashi (1982), Summers (1982) and DeLong and Summers (1986).

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interactions with labor supply (Mankiw, Rotemberg and Summers 1985, Bean 1986, Eichenbaum, Hansen and Singleton 1987); durable goods (Bernanke 1985, Startz 1986); and government purchases (Bailey 1971, Kormendi 1983, Aschauer 1985, Bean 1986). We examine whether these alternative formulations of preferences can explain the apparent excess sensitivity of consumption to income.

The organization of the paper is as follows. Section 2 describes our model and instrumental variables test procedure in more detail, and relates our approach to the existing literature. Section 3 reports empirical results for the basic model. Section 4 presents some Monte Carlo results to shed light on the finite sample properties of our tests. Section 5 extends the model to allow for the effects of time-varying real interest rates and nonseparabilities in the utility function. Section 6 concludes.

2. An Instrumental Variables Approach to the Permanent Income Hypothesis

Consider an economy in which there are two groups of agents, who receive income Y_{1t} and Y_{2t}. Total income Y_t is just the sum of the income of these two groups: $Y_t = Y_{1t} + Y_{2t}$. We assume that the first group receives a fixed share λ of total income[2], so $Y_{1t} = \lambda Y_t$ and $Y_{2t} = (1-\lambda)Y_t$.

Agents in the first group consume their current income, so $C_{1t} = Y_{1t}$. Taking first differences, $\Delta C_{1t} = \Delta Y_{1t} = \lambda \Delta Y_t$[3]. Agents in the second group, by contrast, consume their permanent income: $C_{2t} = Y^P_{2t} = (1-\lambda)Y^P_t$. By the argument of Hall (1978), as elaborated in Flavin (1981), we then have $\Delta C_{2t} = \mu + (1-\lambda)\epsilon_t$, where μ is a constant and ϵ_t is the innovation between time t-1 and time t in agents' assessment of total permanent income Y^P_t. Since ϵ_t is an innovation, it is orthogonal to any variable which is in agents' information set at time t-1.

The change in aggregate consumption can now be written as

$$(1) \quad \Delta C_t = \Delta C_{1t} + \Delta C_{2t} = \mu + \lambda \Delta Y_t + (1-\lambda)\epsilon_t.$$

Our empirical strategy will be to estimate λ, and test the permanent income hypothesis that $\lambda=0$, by running the regression (1). It is

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[2] Y is defined to include both labor and capital income. One might argue that group one should receive a constant share of labor income, rather than total income. In practice, this difference is probably not important.

[3] A slightly more general specification would be that agents in the first group consume a fraction k of their current income. We would then have $k\lambda$ wherever λ appears in the equations below. Our estimates of the income share of current-income consumers are biased downwards if these agents have a marginal propensity to consume less than unity.

important to note, however, that (1) cannot be estimated by Ordinary Least Squares. The error term ϵ_t is orthogonal to lagged variables, but not necessarily to ΔY_t. In fact, most plausible income generating processes will lead to a positive correlation between the change in current income, ΔY_t, and the revision in agents' forecasts of future income, ϵ_t. The correlation will make OLS estimates of λ inconsistent and biased upward[4].

The solution to this problem is to estimate (1) by instrumental variables rather than OLS. Any lagged stationary variables are potentially valid instruments since they are orthogonal to ϵ_t if the model is correct. Of course, good instruments must also be correlated with ΔY_t. If ΔY_t is completely unpredictable then there are no instruments which are orthogonal to ϵ_t but correlated with ΔY_t, and the procedure breaks down. In this case permanent income and current income are equal so the parameter λ is unidentified. More generally, if ΔY_t is only slightly predictable it will be hard to obtain a precise estimate of the parameter λ.

Equation (1), estimated by instrumental variables, can be thought of as a restricted version of a more general two-equation system in which ΔC_t and ΔY_t are regressed directly on the instruments. If we have K instruments, X_{1t} through X_{Kt}, then the general system is

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[4] One can write down special cases in which ΔY and ϵ are uncorrelated. For example, if ΔY is white noise, but consumers know the realization of ΔY perfectly one period in advance, then ΔY and ϵ are uncorrelated. There may also be specific shocks -- such as a deficit-financed increase in military spending, discussed further below -- which raise current disposable income but lower permanent disposable income. Nevertheless, it seems reasonable to assume that an increase in current income is generally associated with an increase in permanent income.

$$(2) \quad \Delta C_t = \beta_0 + \beta_1 X_{1t} + \ldots + \beta_K X_{Kt} + \eta_{Ct} = X_t \beta + \eta_{Ct}$$

$$\Delta Y_t = \gamma_0 + \gamma_1 X_{1t} + \ldots + \gamma_K X_{Kt} + \eta_{Yt} = X_t \gamma + \eta_{Yt}.$$

The permanent income hypothesis implies that the vector $\beta = 0$ (that is, $\beta_1 = \ldots = \beta_K = 0$). This can be tested directly, and without any need for predictability of ΔY_t, by OLS estimation of the first equation of (2). But it is hard to interpret a rejection of the permanent income hypothesis in this framework; an estimate of λ is much more informative about the economic importance of deviations from the theory. For this reason we focus on instrumental variables estimation of (1).

When there is more than a single instrument, equation (1) places over-identifying restrictions on (2), that predictable changes in consumption and income, and therefore the vectors β and γ, are proportional to one another ($\beta = \lambda\gamma$, or $\beta_1/\gamma_1 = \ldots = \beta_K/\gamma_K = \lambda$). If we are to put much weight on the estimate of λ that we obtain from (1), it is important to test these restrictions. A simple way to do this is to regress the residual from the instrumental variables regression on the instruments, and then to compare T times the R^2 from this regression, where T is the sample size, with the χ^2 distribution with (K-1) degrees of freedom. We use this test below.

Equation (1) also implies that for any value of λ, the R^2 of the regression of ΔC_t on instruments must be less than the R^2 of the regression of ΔY_t on instruments, unless ϵ_t and ΔY_t are strongly

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negatively correlated[5]. As argued above, there is a presumption that ϵ_t and ΔY_t are positively correlated, so we expect to find this ordering of the R^2 statistics when we estimate the two-equation system (2) by OLS. This reasoning implies that a small R^2 for changes in consumption cannot be interpreted as strong evidence in favor of the permanent income hypothesis. If the R^2 for changes in income is small, it is very possible that consumption is close to a random walk as measured by R^2 but the permanent income hypothesis is far from true as measured by λ.

The choice of instruments is critically important in our approach. Perhaps the most obvious instruments are ones which summarize the history of Y_t. Flavin (1981) used lagged values of detrended Y_t in her test of the model[6]. Mankiw and Shapiro (1985), however, showed that this leads to statistical problems when the Y_t process has a unit root. Lagged values of ΔY_t are valid instruments but, as we show below, they do not explain a large fraction of the variance of ΔY_t.

Campbell (1987) emphasizes that the history of C_t should also provide good instruments for ΔY_t. This is because, according to the permanent income hypothesis, C_t summarizes agents' information about the future of the Y_t process. If agents have better information about Y_t than is contained in that variable's own history, then C_t will help to forecast

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[5] To see this, note that the R^2 in the consumption equation is $\lambda^2 \text{Var}(X_t\gamma)/(\lambda^2\text{Var}(\Delta Y_t)+(1-\lambda)^2\text{Var}(\epsilon_t)+2\lambda(1-\lambda)\text{Cov}(\Delta Y_t,\epsilon_t))$ which is less than or equal to the R^2 in the income equation when $(1-\lambda)^2\text{Var}(\epsilon_t)+2\lambda(1-\lambda)\text{Cov}(\Delta Y_t,\epsilon_t) \geq 0$.

[6] She estimated the system (2) by OLS, and tested the zero restrictions on the coefficients in the consumption equation. However she motivated and interpreted her results using the model (1).

Y_t. Furthermore, the permanent income hypothesis implies that C_t and Y_t are cointegrated so that saving $S_t = Y_t - C_t$ is stationary. Lagged values of S_t or ΔC_t are likely to increase the precision with which the parameter λ can be estimated.

There is another advantage to using lags of ΔY_t, ΔC_t and S_t as instruments[7]. The unrestricted system (2) which results from this choice of instruments is an error-correction model for consumption and income, of the type proposed by Davidson, Hendry, Srba, and Yeo (1978) and Davidson and Hendry (1981). An error-correction model is an appealing way to summarize the time series behavior of cointegrated variables. Davidson et al. interpret their error-correction models in terms of disequilibrium adjustment of consumption to income; our approach suggests an alternative interpretation, involving forward-looking consumption behavior of at least some agents. As discussed above, our model (1) places testable restrictions on the error-correction framework.

Financial variables are also appealing instruments. There is considerable evidence that changes in stock prices and interest rates help to forecast changes in income (Fischer and Merton 1984, Sims 1980, Litterman and Weiss 1985). Hall (1978) found that stock prices also forecast changes in consumption. We use both stock prices and interest rates in our empirical work.

To conclude this section, we briefly argue that our specification (1) is consistent with the conclusions reached in some other recent work on aggregate consumption behavior. First, we develop the implications of

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[7] Note that $\Delta S = \Delta Y - \Delta C$, so there is potentially a linear dependence among the instruments. This can be avoided by including only one lag of S, and any number of lags of ΔY and ΔC.

(1) for savings behavior and the variability of consumption. According to our model, aggregate saving is given by

$$(3) \quad S_t = Y_t - C_{1t} - C_{2t} = (1-\lambda)(Y_t - Y_t^P).$$

When the permanent income hypothesis holds (that is, when $\lambda = 0$), then saving equals transitory income, the difference between current and permanent income. As λ increases, consumption moves more closely with current income and saving becomes a smaller fraction of transitory income. Thus the model (1) predicts that observed saving is perfectly correlated with its value under the permanent income hypothesis, but that it may have a smaller standard deviation. This implication accords quite well with the results of Campbell (1987).

The effect of λ on the smoothness of consumption can be read off from equation (1). As λ increases, ΔY_t gets greater weight in the consumption change, and ϵ_t gets less weight. Consumption becomes smoother if ΔY_t is less variable than ϵ_t. This will be the case if Y_t follows a process which is more persistent than a random walk, in the sense that shocks tend to be amplified rather than damped by the subsequent movements of the series. (An example is $\Delta Y_t = \rho \Delta Y_{t-1} + \epsilon_t$ where $\rho > 0$.) Campbell and Deaton (1987) argue that in postwar U.S. data consumption is smoother than it would be if the permanent income hypothesis were true. The model with $\lambda > 0$, in which consumption moves too closely with current income, is one possible explanation for this result[8].

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[8] However the model with $\lambda > 0$ cannot explain the observation that consumption is smoother even than current income. Other considerations (such as the arguments for moving average behavior of consumption given

Our model (1) may also be able to account for the finding of Hall (1986) that the response of consumption to contemporaneous changes in military spending is almost zero, even though current income responds positively to these changes. If increased military spending is initially financed by borrowing, with higher taxes coming later, it will raise current disposable income but lower permanent disposable income. In our notation, ΔY_t will be positive but ϵ_t will be negative. The positive response of current-income consumers to ΔY_t could be just offset by the negative response of permanent-income consumers to ϵ_t. Hall (1986) is unable to distinguish these effects because he uses a contemporaneous instrument rather than lagged instruments to estimate the response of consumption to income.

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in the next section) are required to explain this. Deaton (1986) proposes a habit-formation model (his equation (22)) which is our equation (1) with ΔY_t replaced by ΔC_{t-1}.

3. Empirical Results for the Basic Model

Before we can estimate our model, we need to address two issues of specification which arise from the nature of the aggregate time series on consumption and income.

Our discussion so far has been couched in terms of levels and differences of the raw series C_t and Y_t. This is appropriate if these series follow homoskedastic linear processes in levels, with or without unit roots. In fact, however, aggregate time series on consumption and income appear to be closer to log-linear than linear. The mean change and the innovation variance both grow with the level of the series. A correction of some sort appears necessary[9].

Two alternative strategies are available for scaling the variables. One approach is simply to take logs of all the variables in the previous section. Equation (1) should hold in logs, with $\lambda=0$, if aggregate consumption is chosen by a representative agent with a power utility function facing a constant riskless real interest rate (Hansen and Singleton 1983, Bean 1986, Hall 1987, Nelson 1987). The instruments discussed in the previous section remain stationary, but we now use the difference between log consumption and log income, the log consumption-income ratio, rather than saving. The only problem with this approach is that the parameter λ can no longer be precisely interpreted as the fraction of agents who consume their current income; however, if one is willing to approximate the log of an average by an average of logs, the interpretation of the model is not substantially affected.

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[9] Hall (1978) did not scale his variables. Nelson (1987) shows that this has some effect on the results he obtained, although Hall's general conclusions are not sensitive to scaling.

An alternative scaling method is to divide ΔC_t and ΔY_t by the lagged level of income, Y_{t-1}. Campbell and Deaton (1987) derive a linear approximation to the permanent income model which uses this approach. The instruments can be scaled in the same way. In practice both scaling methods give very similar results, so we report only results for logs below[10]. We use lower-case letters to denote log variables.

A second data problem is that consumption and income are measured as quarterly averages rather than at points in time. If the permanent income hypothesis holds in continuous time, then measured consumption is the time average of a random walk. The results of Working (1960) imply that it will have a first-order serial correlation of 0.25, which could lead us to reject the model even if it is true. Christiano, Eichenbaum and Marshall (1987) and Hall (1987) argue that a continuous-time version of the permanent income model fits postwar U.S. data better than a discrete-time version.

We deal with this problem by lagging the instruments more than one period, so that there is at least a two-period time gap between the instruments and the variables in equation (1). The time average of a continuous-time random walk is uncorrelated with all variables lagged more than one period, so by using twice-lagged instruments we obtain a

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[10] The main difference is that the second method gives slightly smaller (but no less statistically significant) estimates of the parameter λ. The reason for this is that in the first method, the left hand side of (1) is approximately $\Delta C_t / C_{t-1}$, while in the second method it is $\Delta C_t / Y_{t-1}$. C_{t-1} is smaller than Y_{t-1}, both because the mean of saving is positive and because we use only nondurables and services consumption, a fraction of the total, in our tests.

test of the model that is valid for time-averaged data[11].

The extra lag in the instruments also helps meet several other potential objections. First, Goodfriend (1986) has noted that aggregate variables are not in individuals' information sets contemporaneously because of delays in government publication of aggregate statistics. Since such delays are typically no more than a few months, lagging the instruments an extra quarter largely avoids this problem[12]. Second, it is sometimes suggested that those goods labelled non-durable in the National Income Accounts are in fact partly durable. Durability would introduce a first-order moving average term into the change in consumer expenditure (Mankiw 1982); this would not affect our procedure using twice-lagged instruments. Third, there may be white noise errors in the levels of our consumption and income variables. These could be due to "transitory consumption", or to measurement errors. White noise errors in levels become first-order moving average errors in our differenced specification, and could be correlated with once-lagged instruments; but they cannot be correlated with twice-lagged instruments.

These arguments for twice-lagging our instruments also imply that the error terms in equations (1) and (2) have a first-order moving average structure. If we ignore this and use standard OLS and instrumental variables procedures, the coefficient estimates remain consistent but the

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[11] Another response to the time-averaging problem would be to use monthly data at quarterly intervals. Nelson (1987) uses this approach. However we found that the results we obtained were somewhat sensitive to whether we used first-month, middle-month or last-month data from each quarter.

[12] The problem is not completely avoided, since the data are revised over a long period of time. Below we use as instruments financial variables such as nominal interest rates, which are known contemporaneously, and this fully circumvents the problem.

standard errors are inconsistent. Fortunately, a straightforward

standard error correction is available (White 1984); White's methods can

also be used to allow for conditional heteroskedasticity in the error

terms of (1) and (2). For our data, these corrections make almost no

difference and we report uncorrected standard errors below[13].

To estimate our model, we use standard U.S. quarterly time series

data, obtained from the Data Resources, Inc. data bank. Y_t is measured

as disposable income per capita, in 1982 dollars. C_t is per capita

consumption of nondurables and services, in 1982 dollars. The data are

available from 1948:1 through 1986:4, but we end our sample in 1985:4 in

order to avoid using highly preliminary 1986 data. We begin our sample

either in 1953:1, the date used by Blinder and Deaton (1985), Campbell

(1987) and Campbell and Deaton (1987), which avoids the Korean War, or in

1949:1, the earliest date which allows us to include instruments lagged

up to two years. The 1949:1 starting date corresponds more closely to

the empirical work of Hall (1978), Flavin (1981) and Nelson (1987).

Table 1 reports results for the 1953:1-1985:4 sample period. The

table has six columns. The first gives the row number and the second the

instruments used[14]. The third and fourth columns give the adjusted R^2

statistics for OLS regressions of Δc_t and Δy_t, respectively, on the

instruments. In parentheses we report the p-value for a Wald test of the

hypothesis that all coefficients are zero except the intercept. The

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[13] Taking account of the moving average error structure tends to
reduce the reported standard errors very slightly; taking account of
heteroskedasticity tends to increase them very slightly.

[14] A constant term is always included as both an instrument and a
regressor, but is not reported in the tables.

fifth column gives the instrumental variables estimate of λ, with an asymptotic standard error. The final column gives the adjusted R^2 statistic for an OLS regression of the residual from the instrumental variables regression on the instruments. In parentheses we report the p-value for the corresponding test of the overidentifying restrictions placed by equation (1) on the general system (2). This test can only be carried out when more than one instrument is used.

The first row of Table 1 shows the coefficient obtained when we estimate equation (1) by OLS. The coefficient is about 0.33 (and the R^2, not reported in the table, is also about 0.33, which means that the variance of consumption growth is one third the variance of income growth). The remaining rows give instrumental variables results for various choices of instruments. In all cases we include at least lags two through four of the instruments[15]; in some rows we add lags five and six, for a total of five instruments.

Rows 2 and 3 of the table use lagged income growth rates as instruments. These are not strongly jointly significant in predicting consumption or income growth; in row 3, for example, lags two through six of income growth are jointly significant at the 16.5% level for consumption growth and at the 6.2% level for income growth. Nevertheless, we estimate λ at 0.477 with an asymptotic standard error of 0.151 in this row. The corresponding t statistic is 3.15, with a

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[15] When we included only the second lag, we found that it was not possible to forecast either consumption or income growth at conventional significance levels.

significance level of 0.2%[16]. The instrumental variables procedure therefore rejects the permanent income hypothesis much more strongly than the OLS test for joint significance of the instruments in predicting consumption growth. This pattern is found consistently throughout the table.

Stronger results are obtained in rows 4 and 5 of the table, where we use lagged consumption growth rates as instruments. It is striking that lagged consumption forecasts income growth more strongly than lagged income itself does, and this enables us to estimate the parameter λ more precisely. It is estimated at 0.526 in row 5 (with a t of 4.17, significant at less than the 0.1% level). The OLS test also rejects the permanent income model in row 5.

We next consider using some financial variables as instruments. We tried using lagged changes in real stock prices (the quarterly percentage change in the real value of the Dow Jones Industrial Average), but found that this variable had no predictive power for consumption growth or income growth[17]. Results using lagged changes in quarterly average 3-month nominal Treasury bill rates (i_t) were much more successful, and we

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[16] Close inspection of the regressions underlying row 3 shows that the fifth lag of income growth predicts both consumption growth and income growth (with a t statistic of -2.25 for consumption, and -2.20 for income). This fact presumably accounts for the strong rejection of the permanent income hypothesis in row 3, as contrasted with the weak evidence in row 2. Campbell (1987) also rejected the permanent income hypothesis more strongly when he included five lags of income growth.

[17] This finding contrasts with the positive results for stock prices reported by Hall (1978) and others. However close inspection of Hall's stock price regression (his equation (8), on p.984) suggests that almost all the explanatory power comes from the first lagged stock price change. When we include the first lag, we also find strong predictive power from stock price changes; but for the reasons discussed above, we regard this as an illegitimate test of the permanent income model.

report these in rows 6 and 7 of Table 1. The instruments are jointly significant for consumption growth at the 1.2% and 0.1% levels. The parameter λ is estimated at 0.713 in row 6 (with a t of 2.94, significant at the 0.4% level), and at 0.615 in row 7 (with a t of 4.49, significant at less than the 0.1% level).

The last two rows of the table report restricted error-correction models for consumption and income. Row 8 has lags of consumption growth, income growth and the log consumption-income ratio as instruments; row 9 adds lagged interest rate changes. The results are broadly consistent with those in earlier rows.

Table 1 also tests the overidentifying restrictions of our model (1) on the unrestricted system (2). The test results are reported in the last column of the table; there is no evidence against our restrictions anywhere in this column[18].

While this is reassuring, we should note that some other features of the results are puzzling. In laying out our model, we argued that one would expect a positive correlation between Δy_t and ϵ_t; this would bias upwards the OLS estimates of λ, and would give a smaller R^2 for the regression of consumption growth on instruments than for the regression of income growth on instruments. In fact, in Table 1 we find that our instrumental variables procedure always estimates λ to be _larger_ than the OLS estimate of 0.328, and in three out of eight cases we find that the

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[18] Deaton's (1986) model, which is our equation (1) with once lagged consumption growth replacing contemporaneous income growth, would cause us to reject the restrictions of our model when variables predicting once lagged consumption growth are included in the instrument set. We regard these results as preliminary evidence that Deaton's model does not explain our findings.

adjusted R^2 in the consumption equation is larger than the adjusted R^2 in the income equation. Consumption growth is surprisingly predictable, given the predictability of income growth.

One possible explanation for this pattern of results is that some of the factors which require us to lag our instruments twice, also reverse the presumption that Δy_t and ϵ_t are positively correlated. For example measurement error in income, uncorrelated with our instruments, could bias downward the OLS estimate of λ and reduce the R^2 in a regression of income growth on instruments, while not affecting the instrumental variables estimates of λ. We regard this explanation as a tentative one, however.

Summarizing table 1, we have found strong evidence against the permanent income hypothesis. The results from our instrumental variables test are particularly unfavorable to the permanent income model. When we use instruments which are jointly significant for predicting income growth at the 5% level or better, we get estimates of λ, the fraction of the population which consumes its current income, in the range 0.35 to 0.65. These estimates are always strongly significant even though we have lagged the instruments two periods instead of one. The overidentifying restrictions of our model are not rejected at any reasonable significance level.

In Table 2A we extend our sample period backwards to 1949. It turns out that the addition of the Korean War to the sample has a powerful effect on our results. The fifth and sixth lags of income now have strong predictive power for consumption growth (row 3), but the coefficients do not obey the restrictions of our model (1); the parameter

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λ is estimated small, negative and insignificant, and the test of our model in the last column rejects at the 0.3% level. Lagged interest rates give estimates of λ which are comparable to those in Table 1, but in general there is much less evidence in 1949-85 that predictable consumption growth is associated with predictable income growth.

It is remarkable that adding only four years of data can have such a large effect on our results. Upon inspecting the data, we found two very unusual observations for income growth. Disposable income grew 6.5% (26% at an annualized rate) in 1950:1, and 4.6% (18% at an annualized rate) in 1975:2. The latter episode was due to a temporary tax rebate (Blinder and Deaton 1985).

In Table 2B we report the results of a simple experiment to see whether our 1949-85 results are dominated by the first quarter of 1950. We set consumption and income growth for that quarter equal to their means for the 1949-85 period, and repeat Table 2A[19]. The results are reassuring. The ability of the fifth and sixth lags of income growth to predict consumption growth is greatly reduced, the point estimates of λ are much closer to those in Table 1 and are often statistically significant, and the model (1) is rejected at the 5% level only in row 5.

We also checked that the Table 1 results are not dominated by the data from 1975:2. Setting the 1975:2 observations to their 1953-85 sample means has only a marginal effect on the Table 1 results. The biggest effect is in row 2, where the estimate of λ falls to 0.264 with a

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[19] This sets the influence of 1950:1 to zero. It is not equivalent to adding a 1950:1 dummy to equation (1), since 1950:1 observations appear in the instrument set as well as on the left and right hand sides of equation (1).

standard error of 0.215. All the other rows are essentially unaffected[20].

As a final check on our results, we split the 1953-85 sample into two even subsamples, 1953:1-1969:2 and 1969:3-1985:4. Results for selected sets of instruments are reported in Table 3. There is a striking contrast between the two subsamples. In the first, income growth is essentially unpredictable using any of our instruments; consequently, it is impossible to identify the parameter λ with any precision or to reject the permanent income hypothesis. In the second subsample, by contrast, we can achieve adjusted R^2 statistics for consumption and income growth of about 25% using five instruments, and the permanent income hypothesis is strongly rejected. Our model (1) is not rejected, indicating that predictable consumption growth is highly correlated with predictable income growth in the 1969-85 period.

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[20] The λ estimates are: row 3 0.465 (0.188); row 4 0.369 (0.142); row 5 0.504 (0.134); row 6 0.747 (0.247); row 7 0.641 (0.143); row 8 0.348 (0.110); row 9 0.452 (0.098).

4. Monte Carlo Results

The evidence in the last section suggests that postwar United States data can reject the permanent income hypothesis. This section presents some Monte Carlo results with two purposes in mind. First, we want to examine the small sample distribution of our test statistics. The problem of small sample bias has been a serious one in tests of the permanent income hypothesis (Mankiw and Shapiro 1985). Monte Carlo experiments can help protect empirical researchers from excessive reliance on asymptotic distribution theory.

Second, we want to shed light on a particular aspect of the results. We found that the hypothesis tests based on the values of λ estimated by instrumental variables imply stronger rejections of the permanent income hypothesis than do the hypothesis tests based on OLS estimation of the unrestricted reduced forms. For example, in Row 3 of Table 1, the random walk of consumption is rejected at only the 16.5% level in the unrestricted reduced form. But the t-statistic on λ is 3.15, indicating a rejection at the 0.2% level. We will reconcile these results by showing that if our alternative hypothesis is correct, the instrumental variables test is more powerful than the unrestricted test.

In Table 4 we report the results of a simple Monte Carlo experiment. We generated 500 data sets, each with 125 observations, from the following process:

(4) $\Delta y_t = u_{1t} + u_{2t} + u_{3t}$

$\Delta c_t = \lambda \Delta y_t + (1-\lambda)u_{1t}.$

21

Here u_{1t}, u_{2t} and u_{3t} are normal random variables, serially uncorrelated and uncorrelated with each other. Current and lagged values of u_{2t} are used as instruments. By choosing $\lambda = 0$, $Var(u_{1t}) = 0.33$, $Var(u_{2t}) = 0.1$, and $Var(u_{3t}) = 1 - 0.33 - 0.1 = 0.57$, we obtain data with the following properties in the population. First, the instruments have no explanatory power for consumption growth. Second, the true coefficient in an OLS regression of Δc_t on Δy_t is 0.33, and the true R^2 of this regression is also 0.33. Finally, the instruments explain 10% of the variation in Δy_t. Our artificial data thus match some basic moments of the actual U.S. data, and satisfy Hall's "random walk" condition[21].

The first panel of Table 4 shows what happens when we apply our methods to these data. In the first row of the panel, we use one instrument, u_{2t}; in the second row, we use three instruments, u_{2t} through $u_{2,t-2}$; and so on up to 15 instruments. Of course, given the data generation process (4), only the first instrument has explanatory power for Δy_t in the population. The table reports the empirical mean estimate of λ across the 500 data sets, and the empirical mean standard error. It also reports the fraction of the data sets for which OLS 5% and 1% tests, and instrumental variables 5% and 1% tests, reject the null hypothesis.

When only one instrument is used, Table 4 shows that the OLS and instrumental variables tests give very similar results. Both reject the null at the 5% level about 5% of the time, and at the 1% level about 1% of the time. The mean instrumental variables estimate of λ is very close

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[21] These data do not satisfy the permanent income hypothesis because the "consumption" process violates the intertemporal budget constraint. For the purpose of evaluating our econometric methods, we believe that this feature of the artificial data is relatively unimportant.

to zero.

As the number of instruments increases, however, the behavior of the instrumental variables test diverges from that of the OLS test. The OLS test tends to reject somewhat less frequently than it should, while the instrumental variables test starts to reject much too frequently. With three instruments, the true size of the instrumental variables test is about twice the theoretical size; with five instruments, the true size of a 5% instrumental variables test is about 15% and the true size of a 1% test is about 5%. With 10 or 15 instruments, the problem of excessive rejection becomes extreme. The mean estimate of λ increases accordingly, and reaches 0.184 when 15 instruments are used.

The reason for this bias in the instrumental variables test is presumably that the first stage regression of Δc_t and Δy_t on instruments tends to "overfit" in finite samples[22]. Since Δc_t and Δy_t are correlated, this tends to give a nonzero coefficient when the fitted value of Δc_t is regressed on the fitted value of Δy_t in the second stage of the instrumental variables procedure. In the extreme case in which one used as many instruments as there are observations, the instrumental variables estimate of λ would be the OLS coefficient of Δc_t on Δy_t, which is about 0.33 in the U.S. data and is exactly 0.33 in the artificial data.

The last column in Part A of Table 4 presents the correct critical values for the IV test. With three instruments, a t-statistic of 2.44 is necessary for a valid test at the 5% level. With five instruments, a

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[22] The same overfitting affects the OLS test, but there it is offset by the increasing degrees of freedom of the test statistic.

critical value of 2.72 is required. These empirical critical values

provide one way to reduce reliance on the asymptotic distribution.

These results imply that one must be careful not to use too many

instruments in testing the permanent income model[23]. But they cannot

explain the strong rejections of the model we obtain in Table 1 using

moderate numbers of instruments. In row 3 of Table 1, for example, the

t-statistic on λ is 3.15, using five instruments. In row 5 the t-

statistic is 4.17 using five instruments. Adjusting for the bias we find

in Table 3 weakens these results somewhat, but the permanent income model

is still easily rejected at the traditional 5% level.

In the second panel of Table 4, we compare the power of the OLS and

instrumental variables procedures against the alternative that λ is

nonzero. We set λ equal to 0.25, keeping the moments of the shocks in

(4) the same as before. We use the empirical critical values for the IV

test. We find that the instrumental variables procedure rejects the

false null hypothesis that λ equals zero much more frequently than the

OLS procedure. The difference is striking even when only one instrument

is used, and it increases with the number of instruments[24]. Using three

instruments, for example, the instrumental variables test rejects at the

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[23] The early literature on instrumental variables estimation
recognized the dangers of using too many instruments; Sargan (1958), for
example, recommended the use of only three instruments in samples of
typical size. But the point seems to have been forgotten in some recent
work testing rational expectations orthogonality restrictions.

[24] It is puzzling that the instrumental variables test tends to
reject more frequently as we increase the number of instruments, even
though the additional instruments have no true explanatory power and we
have corrected the size of the test empirically. This appears to be a
small sample effect; we did not find it present to nearly the same degree
when we ran a small Monte Carlo experiment with 500 observations for each
run.

1% level 32% of the time while the OLS test rejects at this level only 5% of the time. We conclude that the instrumental variables test rejects more strongly than the OLS test in Table 1 primarily because of its greater power to detect this kind of deviation from the permanent income model.

5. <u>Generalizations of the Permanent Income Hypothesis</u>

Our evidence in Section 3 suggests that the aggregate data are not consistent with the permanent income hypothesis. Moreover, the departure from the hypothesis is substantively large: approximately 40% or 50% of income accrues to individuals who consume their current income rather than their permanent income. The purpose of this section is to examine whether generalizations of the permanent income hypothesis along some dimension can explain these findings.

5.1 <u>Changes in the Real Interest Rate</u>

Hall's (1978) "random walk" theorem for consumption rests on the crucial assumption that the real interest rate is constant. Any rejection of the theory might be attributable to the failure of this assumption. For example, Michener (1984) shows how variation through time in the real interest rate can make consumption appear excessively sensitive to income, even though individuals intertemporally optimize in the absence of borrowing constraints. It is therefore important to examine whether the departure from the theory documented above is an artifact of the assumed constancy of the real interest rate.

The generalization of the consumer's Euler equation to allow for changes in the real interest rate is now well-known (Grossman and Shiller 1981, Mankiw 1981, Hansen and Singleton 1983, Hall 1987). The log-linear version of the Euler equation is

$$(5) \quad \Delta c_t = \mu + (1/\alpha)r_t + \epsilon_t,$$

where r_t is the real interest rate contemporaneous with Δc_t, and as

before the error term ϵ_t may be correlated with r_t but is uncorrelated with lagged variables. According to (5), high ex ante real interest rates should be associated with rapid growth of consumption. If higher income growth is associated with higher real interest rates, the deviation from the permanent income hypothesis documented above could be explained by this more general model.

To examine this possibility, we consider a more general model in which a fraction λ of income goes to individuals who consume their current income and the remainder goes to individuals who satisfy the general Euler equation (5). We estimate by instrumental variables

$$(6) \quad \Delta c_t = \mu + \lambda \Delta y_t + \theta r_t + \epsilon_t,$$

where $\theta = (1-\lambda)/\alpha$. We thus include the actual income growth and the ex post real interest rate in the equation, but instrument using twice lagged variables. The nominal interest rate we use is the average three-month treasury bill rate over the quarter, the price index is the deflator for consumer nondurables and services, and we assume that there is a 30% marginal tax rate on interest[25]. The results are in Table 5.

We find no evidence that the ex ante real interest rate is associated with the growth rate of consumption. The coefficient on the real interest rate is consistently less than its standard error. Moreover, the coefficient on current income remains substantively and statistically significant. In contrast to the suggestion of Michener (1984), the

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[25] We obtained similar results when we assumed a marginal tax rate of zero.

excess sensitivity of consumption to income cannot be explained by fluctuations in the real interest rate.

The regressions of Table 5 are similar to those of Hall (1987), except that they include the change in income as well as the ex post real interest rate. When we omitted the change in income and estimated equation (5), we found slightly larger, but still fairly small values for the coefficient on the real interest rate: 0.202 (with a standard error of 0.115) for the instruments in row 2 of Table 5, 0.140 (0.117) for the instruments in row 3 of Table 5, and 0.180 (0.119) for the instruments in row 4 of Table 5.

Hall interprets evidence of this sort as indicating that consumers are extremely reluctant to substitute intertemporally ($1/\alpha$ is very small, and α is very large). We note, however, that if equation (5) is properly specified one should be able to estimate it with either consumption growth or the real interest rate as the dependent variable. When the real interest rate is the dependent variable, the coefficient estimated by instrumental variables should be α rather than ($1/\alpha$). On reversing the regressions, we found coefficients of 1.469 (0.503), 0.825 (0.337), and 0.772 (0.287) respectively. While these estimates are statistically significant, and larger than those from the Hall regressions, one cannot conclude from them that α is very large.

The reason why we obtain such different results when we renormalize equation (5) is that the data reject the overidentifying restrictions of (5). The reversed regressions are all rejected at the 0.1% level or better, indicating that there are predictable movements in real interest

rates which are not associated with predictable consumption growth[26].

One should be cautious in interpreting estimates from such a system. Our

model (1), however, is not sensitive to renormalization and we do not

reject its restrictions in either its original or its renormalized form.

5.2 Nonseparabilities in the Utility Function

The "random walk" theorem for consumption will also fail if

consumption is not separable in the utility function from other goods.

With constant real interest rates, the marginal utility of consumption is

a martingale even under non-separability. That is, it is still true that

$$(7) \quad E_t U'(C_{t+1}, X_{t+1}) = \gamma U'(C_t, X_t)$$

for some constant γ. Yet predictable changes in the other good X must

lead to predictable changes in consumption to maintain the martingale

property of marginal utility. If changes in X are correlated with

changes in income, non-separability could in principle explain the

apparent excess sensitivity of consumption to income documented in

Section 3.

We test for non-separability in a very simple way. We include the

change in log X as an additional right-hand side variable in our

equation. This functional form can be formally justified if the utility

function is Cobb-Douglas (Bean 1986) or as a log linear approximation to

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[26] Both the specifications (5) and (6) are rejected when the regressions are reversed to make the real interest rate the dependent variable, but are not rejected when the real rate is an explanatory variable. The reason for this is that the real rate is given a coefficient of approximately zero when it is an explanatory variable, so the predictable movements of the real interest rate do not enter into the fitted value or residual of the equation.

a more general specification. As before, we estimate the equation using twice-lagged instrumental variables.

Various non-separabilities have been proposed. Mankiw, Rotemberg, and Summers (1985) and Eichenbaum, Hansen, and Singleton (1987) consider non-separability between consumption and labor supply. In Table 6 we include the change in log labor supply as a right hand side variable with coefficient θ. Labor supply is measured as per capita man-hours in nonagricultural establishments. The results suggest no important non-separability between consumption and labor supply. Even though there is substantial predictable variation in the quantity of labor supplied, it apparently does not lead to predictable changes in consumption.

Bernanke (1985) and Startz (1987) propose that the marginal utility of nondurable goods may be affected by the stock of consumer durable goods. In Table 7 we enter this stock as the X variable. The stock of consumer durables during a quarter is measured as the average of the end-of-quarter stock and the previous end-of-quarter stock[27]. (Other timing assumptions lead to similar results.) We find substantial predictable changes in the stock of durables, but no evidence that these changes coincide with predictable changes in consumption.

It is often suggested that changes in government purchases of goods and services affect the marginal utility of private consumption (Bailey 1971, Kormendi 1983, Aschauer 1985). Indeed, Aschauer suggests that allowing for such an effect can save the consumption Euler equation from a statistical rejection. In Table 8 we examine this possibility by

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[27] We constructed an end-of-quarter stock series from the annual stock at the beginning of the sample period and the series on consumer durable purchases, assuming a depreciation rate of 6% per quarter.

entering the change in the log of total government purchases per capita as a right-hand side variable. Again, we find no evidence of non-separability in the utility function. Moreover, the estimate of λ we obtain remains statistically and substantively significant. In contrast to Aschauer, we find that non-separability between private and public purchases does not improve the performance of the consumption Euler equation.

6. Conclusions

Our analysis of United States postwar quarterly data leads us to the following conclusions:

(1) There is modest evidence against the implication of the permanent income hypothesis that changes in consumption are unforecastable. When the change in log consumption is regressed on its own lags two through six in the 1953-85 period, the null hypothesis that all the coefficients are zero can be rejected at the 0.6% level. While the adjusted R^2 of this regression is small (9%), a small R^2 should not be viewed as supportive of the permanent income hypothesis, since the R^2 of the comparable regression for the change in disposable income is also very small.

(2) The evidence against the permanent income model comes primarily from the second half of our sample period, 1969-85. In the first half of the sample, 1953-69, the data have little power to discriminate between models because income growth is essentially unpredictable.

(3) The forecastability of consumption can be explained by a model in which a fraction λ of income goes to individuals who consume their current income rather than their permanent income. This more general model is not statistically rejected. Our estimates suggest that λ is approximately 0.4 or 0.5, indicating a substantial departure from the permanent income hypothesis.

(4) The result that consumption tracks income too closely cannot be explained by the time-averaged nature of the data, by short delays in publication of aggregate statistics, or by partial durability of goods labelled "nondurable" in the National Income Accounts. Our test of the

permanent income model is robust to all these problems because, in common with Hall (1987) but in contrast with much of the rest of the literature, we lag our instruments by two quarters instead of one.

(5) Our results cannot be explained by appealing to more general versions of the permanent income hypothesis. We have allowed for changes in the real interest rate, but we find no evidence that changes in the real interest rate lead to predictable changes in consumption. We have also allowed for non-separability in the utility function between consumption and other goods--labor supply, consumer durables, and government purchases--but we find no evidence for any such non-separability.

TABLE 1

BASIC MODEL, 1953-85

Row	Instruments	First-stage regressions Δc equation	Δy equation	λ estimate (s.e.)	Test of restrictions
1	None (OLS)	-----	-----	0.328 (0.041)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}$	-0.003 (0.460)	0.021 (0.124)	0.379 (0.196)	-0.023 (0.988)
3	$\Delta y_{t-2}, \ldots, \Delta y_{t-6}$	0.022 (0.165)	0.043 (0.062)	0.477 (0.151)	-0.035 (0.962)
4	$\Delta c_{t-2}, \ldots, \Delta c_{t-4}$	0.022 (0.121)	0.050 (0.024)	0.406 (0.154)	-0.012 (0.486)
5	$\Delta c_{t-2}, \ldots, \Delta c_{t-6}$	0.089 (0.006)	0.088 (0.006)	0.526 (0.126)	-0.016 (0.555)
6	$\Delta i_{t-2}, \ldots, \Delta i_{t-4}$	0.062 (0.012)	0.025 (0.099)	0.713 (0.242)	-0.016 (0.601)
7	$\Delta i_{t-2}, \ldots, \Delta i_{t-6}$	0.122 (0.001)	0.087 (0.007)	0.615 (0.137)	-0.025 (0.771)
8	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}, \Delta c_{t-2}, \ldots, \Delta c_{t-4}, c_{t-2} - y_{t-2}$	0.010 (0.305)	0.089 (0.011)	0.351 (0.110)	-0.040 (0.914)
9	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}, \Delta c_{t-2}, \ldots, \Delta c_{t-4}, \Delta i_{t-2}, \ldots, \Delta i_{t-4}, c_{t-2} - y_{t-2}$	0.080 (0.031)	0.115 (0.007)	0.454 (0.099)	-0.031 (0.715)

Notes: The statistics in columns 3, 4 and 6 are adjusted R^2 values, and significance levels for tests of the hypothesis that all coefficients except the constant are zero (in parentheses). The statistics in column 5 are the instrumental variables estimate of λ, with an asymptotic standard error (in parentheses).

BASIC MODEL, 1949-85

Row	Instruments	First-stage regressions Δc equation	Δy equation	λ estimate (s.e.)	Test of restrictions
1	None (OLS)	-----	-----	0.217 (0.040)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}$	-0.009 (0.639)	0.055 (0.012)	-0.095 (0.176)	-0.013 (0.565)
3	$\Delta y_{t-2}, \ldots, \Delta y_{t-6}$	0.083 (0.005)	0.042 (0.050)	-0.033 (0.166)	0.077 (0.003)
4	$\Delta c_{t-2}, \ldots, \Delta c_{t-4}$	0.004 (0.307)	0.018 (0.130)	0.153 (0.208)	0.005 (0.157)
5	$\Delta c_{t-2}, \ldots, \Delta c_{t-6}$	0.007 (0.309)	0.011 (0.247)	0.253 (0.191)	0.002 (0.255)
6	$\Delta i_{t-2}, \ldots, \Delta i_{t-4}$	0.045 (0.022)	0.014 (0.170)	0.694 (0.306)	-0.016 (0.712)
7	$\Delta i_{t-2}, \ldots, \Delta i_{t-6}$	0.088 (0.003)	0.053 (0.027)	0.604 (0.176)	-0.027 (0.888)
8	$\Delta y_{t-2}, \ldots, \Delta y_{t-4},$ $\Delta c_{t-2}, \ldots, \Delta c_{t-4},$ $c_{t-2} - y_{t-2}$	0.003 (0.387)	0.149 (0.000)	0.074 (0.097)	0.004 (0.270)
9	$\Delta y_{t-2}, \ldots, \Delta y_{t-4},$ $\Delta c_{t-2}, \ldots, \Delta c_{t-4},$ $\Delta i_{t-2}, \ldots, \Delta i_{t-4}$ $c_{t-2} - y_{t-2}$	0.047 (0.086)	0.153 (0.001)	0.154 (0.089)	0.046 (0.059)

Notes: See notes to Table 1.

TABLE 2B

BASIC MODEL, 1949-85
1950:1 OBSERVATION REMOVED

Row	Instruments	First-stage regressions Δc equation	Δy equation	λ estimate (s.e.)	Test of restrictions
1	None (OLS)	-----	-----	0.251 (0.045)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}$	-0.014 (0.787)	0.013 (0.174)	0.256 (0.246)	-0.020 (0.913)
3	$\Delta y_{t-2}, \ldots, \Delta y_{t-6}$	0.033 (0.082)	0.006 (0.324)	0.499 (0.250)	0.005 (0.218)
4	$\Delta c_{t-2}, \ldots, \Delta c_{t-4}$	0.012 (0.185)	0.057 (0.010)	0.240 (0.163)	0.004 (0.162)
5	$\Delta c_{t-2}, \ldots, \Delta c_{t-6}$	0.013 (0.225)	0.061 (0.018)	0.288 (0.148)	-0.003 (0.023)
6	$\Delta i_{t-2}, \ldots, \Delta i_{t-4}$	0.045 (0.023)	0.020 (0.117)	0.719 (0.298)	-0.016 (0.683)
7	$\Delta i_{t-2}, \ldots, \Delta i_{t-6}$	0.089 (0.003)	0.073 (0.008)	0.609 (0.166)	-0.025 (0.845)
8	$\Delta y_{t-2}, \ldots, \Delta y_{t-4},$ $\Delta c_{t-2}, \ldots, \Delta c_{t-4},$ $c_{t-2} - y_{t-2}$	-0.007 (0.539)	0.103 (0.003)	0.212 (0.118)	-0.022 (0.678)
9	$\Delta y_{t-2}, \ldots, \Delta y_{t-4},$ $\Delta c_{t-2}, \ldots, \Delta c_{t-4},$ $\Delta i_{t-2}, \ldots, \Delta i_{t-4}$ $c_{t-2} - y_{t-2}$	0.040 (0.113)	0.123 (0.003)	0.327 (0.106)	-0.008 (0.444)

Notes: See notes to Table 1.

TABLE 3

BASIC MODEL, 1953:1-1969:2 AND 1969:3-1985:4

Row	Instruments	First-stage regressions		λ estimate (s.e.)	Test of restrictions
		Δc equation	Δy equation		
A. 1953:1-1969:2					
1	None (OLS)	-----	-----	0.472 (0.072)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-6}$	0.003 (0.386)	-0.061 (0.931)	0.233 (0.551)	0.036 (0.122)
3	$\Delta c_{t-2}, \ldots, \Delta c_{t-6}$	-0.031 (0.673)	-0.059 (0.914)	0.752 (0.535)	-0.046 (0.690)
4	$\Delta i_{t-2}, \ldots, \Delta i_{t-6}$	0.016 (0.301)	0.014 (0.315)	0.638 (0.252)	-0.039 (0.611)
B. 1969:3-1985:4					
5	None (OLS)	-----	-----	0.264 (0.048)	-----
6	$\Delta y_{t-2}, \ldots, \Delta y_{t-6}$	0.077 (0.082)	0.080 (0.076)	0.447 (0.137)	-0.070 (0.936)
7	$\Delta c_{t-2}, \ldots, \Delta c_{t-6}$	0.239 (0.001)	0.237 (0.002)	0.411 (0.095)	0.012 (0.212)
8	$\Delta i_{t-2}, \ldots, \Delta i_{t-6}$	0.235 (0.002)	0.092 (0.058)	0.579 (0.155)	-0.040 (0.619)

Notes: See notes to Table 1.

TABLE 4
MONTE CARLO RESULTS

| Number of instruments | Mean λ estimate (mean std error) | Rejection Probability | | Empirical Crit. Val. |
		OLS 5% test (OLS 1% test)	IV 5% test (IV 1% test)	IV 5% (IV 1%)
A. True λ = 0.00				
1	-0.043 (0.206)	5.0% (0.6%)	5.6% (1.8%)	1.99 (2.87)
3	0.028 (0.170)	6.4% (1.2%)	11.0% (4.0%)	2.44 (3.54)
5	0.071 (0.146)	4.6% (1.0%)	15.4 (7.0)	2.72 (3.58)
10	0.143 (0.118)	3.6% (0.4%)	32.2% (17.8%)	3.34 (4.26)
15	0.184 (0.103)	2.6% (0.6%)	49.6% (27.2%)	3.78 (4.97)
B. True λ = 0.25				
1	0.215 (0.160)	30.0% (10.6%)	47.4% (31.8%)	
3	0.275 (0.122)	20.4% (5.0%)	55.6% (32.0%)	
5	0.309 (0.107)	15.4% (5.2%)	63.2% (40.8%)	
10	0.349 (0.091)	9.2% (3.0%)	69.4% (46.6%)	
15	0.391 (0.077)	8.8% (1.4%)	84.2% (58.6%)	

TABLE 5

REAL INTEREST RATES, 1953-85

Row	Instruments	First-stage regressions			λ (s.e.)	θ (s.e.)	Test of restrictions
		Δc	Δy	r			
1	None (OLS)	-----	-----	-----	0.312 (0.041)	0.136 (0.070)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}$ $r_{t-2}, \ldots, r_{t-4}$	0.031 (0.128)	0.039 (0.191)	0.484 (0.000)	0.413 (0.170)	0.051 (0.116)	-0.015 (0.379)
3	$\Delta c_{t-2}, \ldots, \Delta c_{t-4}$ $r_{t-2}, \ldots, r_{t-4}$	0.046 (0.066)	0.049 (0.057)	0.467 (0.000)	0.451 (0.143)	0.049 (0.106)	-0.010 (0.314)
4	$\Delta i_{t-2}, \ldots, \Delta i_{t-4}$ $r_{t-2}, \ldots, r_{t-4}$	0.077 (0.015)	0.026 (0.155)	0.448 (0.000)	0.668 (0.214)	-0.022 (0.141)	-0.022 (0.515)

TABLE 6

LABOR SUPPLY, 1953-85

Row	Instruments	First-stage regressions			λ (s.e.)	θ (s.e.)	Test of restrictions
		Δc	Δy	Δl			
1	None (OLS)	-----	-----	-----	0.287 (0.044)	0.102 (0.046)	-----
2	$\Delta y_{t-2}, \ldots, \Delta y_{t-4}$ $\Delta l_{t-2}, \ldots, \Delta l_{t-4}$	-0.006 (0.502)	0.025 (0.164)	0.139 (0.001)	0.380 (0.235)	0.037 (0.152)	-0.046 (0.990)
3	$\Delta c_{t-2}, \ldots, \Delta c_{t-4}$ $\Delta l_{t-2}, \ldots, \Delta l_{t-4}$	0.029 (0.138)	0.079 (0.013)	0.221 (0.000)	0.364 (0.203)	0.047 (0.144)	-0.037 (0.841)
4	$\Delta i_{t-2}, \ldots, \Delta i_{t-4}$ $\Delta l_{t-2}, \ldots, \Delta l_{t-4}$	0.086 (0.010)	0.062 (0.031)	0.150 (0.000)	0.497 (0.223)	0.087 (0.172)	-0.028 (0.642)

TABLE 7

DURABLE GOODS, 1953-85

Row	Instruments	First-stage regressions			λ (s.e.)	θ (s.e.)	Test of restrictions
		Δc	Δy	Δd			
1	None (OLS)	-----	-----	-----	0.310 (0.041)	0.148 (0.079)	-----
2	$\Delta y_{t-2},\ldots,\Delta y_{t-4}$ $\Delta d_{t-2},\ldots,\Delta d_{t-4}$	0.021 (0.191)	0.028 (0.141)	0.735 (0.000)	0.290 (0.156)	0.087 (0.093)	0.011 (0.114)
3	$\Delta c_{t-2},\ldots,\Delta c_{t-4}$ $\Delta d_{t-2},\ldots,\Delta d_{t-4}$	0.025 (0.162)	0.054 (0.046)	0.769 (0.000)	0.253 (0.142)	0.097 (0.096)	0.009 (0.129)
4	$\Delta i_{t-2},\ldots,\Delta i_{t-4}$ $\Delta d_{t-2},\ldots,\Delta d_{t-4}$	0.106 (0.003)	0.055 (0.044)	0.752 (0.000)	0.549 (0.155)	0.082 (0.106)	-0.000 (0.200)

TABLE 8

GOVERNMENT SPENDING, 1953-85

Row	Instruments	First-stage regressions			λ (s.e.)	θ (s.e.)	Test of restrictions
		Δc	Δy	Δg			
1	None (OLS)	-----	-----	-----	0.328 (0.041)	-0.006 (0.026)	-----
2	$\Delta y_{t-2},\ldots,\Delta y_{t-4}$ $\Delta g_{t-2},\ldots,\Delta g_{t-4}$	-0.026 (0.843)	0.012 (0.271)	0.067 (0.024)	0.344 (0.202)	0.021 (0.093)	-0.044 (0.979)
3	$\Delta c_{t-2},\ldots,\Delta c_{t-4}$ $\Delta g_{t-2},\ldots,\Delta g_{t-4}$	-0.001 (0.427)	0.043 (0.077)	0.046 (0.065)	0.357 (0.142)	0.051 (0.088)	-0.035 (0.797)
4	$\Delta i_{t-2},\ldots,\Delta i_{t-4}$ $\Delta g_{t-2},\ldots,\Delta g_{t-4}$	0.040 (0.085)	0.013 (0.263)	0.037 (0.097)	0.664 (0.233)	0.103 (0.125)	-0.035 (0.797)

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