Monetary Distortions in the Consumption-Leisure Choice: An Empirical Investigation

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Abstract:

In a wide class of monetary models with both cash and credit goods, the main welfare cost of inflation is that it distorts the choice between these two goods. In these models, distortions exist because the relevant measure of the relative price between cash and credit goods for consumers is the usual relative price discounted by the nominal interest rate. Changes in the inflation rate therefore create distortions by affecting the nominal interest rate. This paper proposes a new statistical method for detecting the existence and magnitude of this distortion in a monetary model of the consumption-leisure choice. The empirical analysis is motivated by deriving a long-run restriction between the stochastic and deterministic trends of real consumption, the real wage rate and the gross nominal interest rate from the first-order conditions of the representative agent's optimization problem. Using nondurable- and food-consumption as cash goods, and leisure as the credit good, this method is applied to a diverse group of 12 economies. The evidence suggests that such distortions exist and tend to be statistically and economically significant for most high- and medium-inflation economies, but not for low-inflation economies.

JEL Classification: **E21 (Consumption), E41 (Demand for Money), C22 (Time-Series Models)**

1 Introduction

In a wide class of monetary models with cash goods (goods purchased with money) and credit goods (goods purchased with credit), the main welfare cost of inflation is that it distorts the choice between these two goods (see, e.g., Lucas (1984), Lucas and Stokey (1987), and Townsend (1987)). This class includes many cash-in-advance models with one good and leisure in which leisure plays the role of a credit good (see, e.g., Schmitt-Grohe and Uribe (2000) and Ireland (2003)). In these models, distortions exist because the relevant measure of the relative price between cash and credit goods for consumers is the usual relative price discounted by the nominal interest rate. Changes in the inflation rate therefore create distortions by affecting the nominal interest rate. This paper proposes a new statistical method for detecting the existence and magnitude of this distortion in a monetary model of the consumption-leisure choice.

In the monetary models cited above, money is held for transactions purposes, and the distortions are caused for the following reason. As long as the nominal interest rate is positive, holding non-interest-bearing money involves an opportunity cost. Consumers count this opportunity cost as an extra cost for purchasing cash goods but not for credit goods because money is held only for transactions involving cash goods.

In our empirical work, we use nondurable- and food-consumption as cash goods and leisure as the credit good. The empirical analysis is motivated by deriving a long-run restriction between the stochastic and deterministic trends of real consumption, the real wage rate and the gross nominal interest rate from the first-order conditions of the representative agent's optimization problem. We investigate a diverse group of 12 economies. The evidence suggests that such distortions exist and tend to be statistically and economically significant for most high- and medium-inflation economies, but not for low-inflation economies.

Empirical work by Hodrick, Kocherlakota, and Lucas (1991) and Braun (1994) also uses monetary models with cash and credit goods, but these authors do not investigate this particular form of distortion. Many other aspects of monetary distortions have been studied in the empirical literature on monetary economics. For example, Fisher (1981) estimates shoe leather costs (the costs in time and effort incurred by people and firms who are trying minimize their holdings of cash). Christiano, Eichenbaum and Evans (1996) try to detect evidence on the empirical plausibility of the "limited participation" models of Christiano and Eichenbaum (1992, 1995), in which monetary distortions occur because some people are not allowed to continuously participate in financial trades. More recently, Boyd, Levine and Smith (2001) examine the evidence in favor of theoretical models (Huybens and Smith, 1999) in which even predictable increases in inflation affect the financial sector's performance adversely due to informational asymmetries in credit markets.

In contrast, monetary distortions on the relative price of cash versus credit goods have only been studied in the theoretical literature, and have not been studied by other researchers in the empirical literature. Ogaki (1988) studied the relative price monetary distortion with U.S. time series data, using food as the cash good and automobiles as the credit good. Because the inflation rate has been relatively low in the United States, he found only mixed evidence of such distortions. *A priori*, it is expected that stronger evidence will emerge from countries with higher inflation rates. Ogaki (1988) used cointegrating regressions that are similar to the ones used here. The concept of cointegration proposed in Engle and Granger (1987) was relatively new in 1988, and better econometric procedures for cointegrated systems have been developed in the last decade. Hence, the econometric procedure used in Ogaki (1988) is not satisfactory for the purpose of our research.

Ogaki and Park (1998) proposed a cointegration approach to estimating preference parameters, which can be readily used for our research. Ogaki and Park's approach has been used by Ogaki (1992), Cooley and Ogaki (1996) and Ogaki and Reinhart (1998), among others. The econometric procedure proposed by Ogaki and Park allows one to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction, both of which are implied by our model. Stochastic cointegration means that the stochastic trends in the variables are eliminated when their linear combination is formed by a vector, called the cointegrating vector. The deterministic cointegration restriction means that the cointegrating vector also eliminates the deterministic trends, which arise from the drift terms of the variables. To test the sensitivity of our empirical results to the econometric procedure, we also use a more recently developed Hausman-type cointegration test proposed in Choi, Hu and Ogaki (2005).

The rest of the paper is organized as follows. Section 2 describes the economic model. Section 3 presents the econometric model based on the model in Section 2. The econometric procedures are explained in Section 4. Section 5 presents the empirical results. Concluding remarks are contained in Section 6.

2 A Cash-In Advance Model of the Consumption-Leisure Choice

Consider the representative consumer who maximizes the lifetime utility function

$$U = \sum_{t=0}^{\infty} \beta^t u(C_t, L_t) \tag{1}$$

subject to appropriate budget constraints and cash-in-advance constraints for purchasing the consumption good. Here β is a discount factor, C_t denotes consumption and L_t denotes leisure at time t. It is assumed that the momentary utility function u is additively separable in consumption and leisure and has the following functional form

$$u(C_t, L_t) = \frac{C_t^{1-\alpha} - 1}{1-\alpha} + V(L_t). \tag{2}$$

Here V is a continuously differentiable concave function and α is a preference parameter that can be interpreted as the inverse of the intertemporal elasticity of substitution of consumption. The first order condition for the consumption-leisure choice can be summarized by

$$\frac{W_t}{(1+i_t)} = \frac{V'(L_t)}{C_t^{-\alpha}} \tag{3}$$

where W_t is the real wage rate, i_t is the nominal interest rate and $V'(L_t)$ is the marginal utility of leisure. Taking the logarithm of both sides of Equation (3) yields

$$\ln(C_t) = \frac{1}{\alpha}\ln(W_t) - \frac{1}{\alpha}\ln(1+i_t) - \ln(V'(L_t)). \tag{4}$$

Equation (4) forms the basis of our econometric model.

The left-hand-side of Equation (3) is the relevant relative price for the consumption-leisure choice because the consumer is required to hold cash in order to purchase the consumption good in this model. Since the opportunity cost of holding cash is the forgone interest payment, Equation (3) is obtained.

3 The Econometric Model

Since the seminal work of Nelson and Plosser (1982), it is well known that most macroeconomic time series are well approximated by unit-root nonstationary processes. Thus, $\ln(C)$, $\ln(W)$, $\ln(1+i)$ are assumed to be unit root nonstationary. This assumption is consistent with the evidence documented in Ogaki (1992) and Cooley and Ogaki (1996).[1] Leisure is assumed to be strictly stationary. This implies that $\ln(V'(L_t))$ is strictly stationary, so Equation (4) gives the cointegrating regression:

$$\ln(C_t) = b_1 + b_2\ln(W_t) + b_3\ln(1+i_t) + u_t \tag{5}$$

[1]The evidence presented in Subsection 5.2 of this paper is also consistent with this assumption.

where $b_1 = E(\ln(V'(L_t)))$, $b_2 = 1/\alpha$, $b_3 = -1/\alpha$, and $u_t = \ln(V'(L_t)) - b_1$. In Equation (5), i_t is the interest rate with the maturity date that exactly matches the holding period of money. However, the holding period of money is not known, and the data for that particular interest rate might not be available even if the holding period of money were known. Therefore, we use available short-term interest rate data for i_t in our cointegrating regression. Using this variable does not violate the cointegration implication of the model as long as the measured interest rate is cointegrated with the interest rate in the model. This assumption is plausible because all interest rates are cointegrated if risk and term premiums are stationary.

The model implies that $b_2 = -b_3 = 1/\alpha$. However, b_3 in our cointegrating regression will be different from $-1/\alpha$ if the measured interest rate is not cointegrated with the interest rate in the model with a $(1, -1)'$ cointegrating vector. For example, the cointegrating vector will be different from $(1, -1)'$ except for the special case that the holding period is exactly equal to one year in our annual data.[2] Hence we do not interpret the estimated b_3 as $1/\alpha$ and do not impose the restriction $b_2 = b_3$ in our empirical work. In addition, if the consumer decides to change the holding period of money as the short-term nominal interest rate changes, then our assumption of the constant holding period of money is violated. Even in this case, Equation (5) as a cointegrating regression may be a good approximation because the short-term nominal interest rate will be a good measure of transaction costs in equilibrium. However, there is no reason to believe that the restriction $b_2 = b_3$ should hold in this case.

4 The Econometric Procedure

If the monetary distortion exists, the model implies that b_3 is negative in Equation (5). Hence, the main purpose of this paper is to see if the data support this prediction of the model. We can apply Ordinary Least Squares (OLS) directly to Equation (5), and this estimator is called static OLS in the cointegration literature. Even though the static OLS estimator is super consistent, it is not attractive for our purpose because it is not asymptotically efficient. It is also difficult to conduct inference on the statistical significance of b_3 with static OLS because it has a nonstandard distribution. Hence we utilize two asymptotically efficient estimators – Park's (1992) Canonical Cointegrating Regression (CCR) estimator[3] and Stock and Watson's (1993) dynamic OLS estimator. The static OLS estimator is not efficient because

[2] For example, if the deviation from the expectation theory of the term structure of interest rates is stationary, then the cointegrating vector is $(1, 1)'$ if the true holding period of money is one month and a one-year interest rate and an annualized one-month interest rates are used.

[3] Monte Carlo experiments in Park and Ogaki (1991) show that they have better small sample properties than Johansen's (1991) estimators even when the Gaussian VAR structure assumed by Johansen is true. Further details regarding CCR-based estimation and testing can be found in Ogaki (1993a, 1993b).

strict exogeneity is violated. The CCR estimator corrects for endogeneity by a nonparametric method. The dynamic OLS estimator corrects for endogeneity by a parametric method of adding leads and lags of the first difference of the regressor. The parametric method would work better if the parametric correction is a good approximation, whereas the nonparametric method is more robust.

We also test for cointegration. Since the model implies cointegration, it is desirable to test the null hypothesis of cointegration to control the probability of rejecting a valid economic model. Although estimation methods that have no cointegration as their null hypothesis are commonly used in the literature, these methods have low power and may fail to reject the null hypothesis with high probability even when the model is actually consistent with the data.

With the CCR procedure, we use Park's (1990) variable addition test to test the null hypothesis of cointegration. With this procedure, we can test the null hypothesis of stochastic cointegration and the null hypothesis of the deterministic cointegration restriction.[4] In the DOLS framework, we use Choi, Hu, and Ogaki's (2005) recently developed Hausman-type cointegration test to test the null hypothesis of cointegration.

While the DOLS estimator is well known, the Hausman-type cointegration test is not. Therefore, the main idea of the Hausman-type cointegration test is explained briefly in this subsection. Consider the following dynamic regression where y_t is an I(1) variable, x_t is a 2-dimensional vector of I(1) variables, and k stands for the order of leads and lags.

$$y_t = \beta' x_t + \sum_{i=1}^{2} \sum_{j=1}^{k} (\gamma_{i,j} \Delta x_{i,t-j}) + e_t. \tag{6}$$

The leads and lags of the first difference of the regressors are added in this regression in order to correct for the endogeneity problem. Following Stock and Watson (1993), we assume that the endogeneity correction of adding leads and lags perfectly eliminates the endogeneity problem in that e_t is strictly exogenous with respect to the regressors in Equation (6) in this paper.[5]

Now, if the error term e_t is I(0), then this is a dynamic cointegrating regression. As shown by Stock and Watson (1993), the OLS estimator for this regression is super-consistent and asymptotically efficient under their regularity conditions.

On the other hand, if the error term e_t is I(1), then Equation (6) is a spurious regression, and the Dynamic OLS is inconsistent for the coefficient β as shown in Choi et al. (2005). The Hausman-type cointegration test utilizes these properties to discriminate between the situation in which e_t is I(0) and

[4]See Ogaki and Park (1998) for definitions of these terminologies and a description of the testing procedure.

[5]It is assumed here that the endogeneity correction is complete only for ease of exposition. This assumption can be relaxed under some regularity conditions as shown by Saikkonen (1991).

that in which e_t is I(1). In order to understand the idea of this test, it is useful to start with an analysis of the DOLS estimator by the Gauss-Markov Theorem, using Ogaki and Choi's (2001) framework. For this purpose, we consider a special case in which e_t is serially uncorrelated. In this case, Ogaki and Choi's conditional probability version of the Gauss-Markov Theorem applies,[6] and the OLS applied to Equation (6) is the Best Linear Unbiased Estimator (BLUE) given the realization of the regressors.

Now consider the case in which e_t is a random walk. Then all the assumptions of the conditional probability version of the Gauss-Markov Theorem hold except for the spherical variance assumption. In this case, the OLS applied to Equation (6) is unbiased (since we are assuming strict exogeneity), but is not efficient. In this case, we can apply the Generalized Least Squares (GLS) to Equation (6) to obtain the BLUE. Applying GLS to Equation (6) basically means that we apply OLS after taking the first difference of Equation (6):

$$\Delta y_t = \beta' \Delta x_t + \sum_{i=1}^{2} \sum_{j=1}^{k} (\gamma_{i,j} \Delta^2 x_{i,t-j}) + \Delta e_t. \tag{7}$$

Choi et al. (2005) call this estimator the GLS corrected estimator.

In more general cases in which the error is serially correlated, Choi et al. (2005) note that asymptotic theory shows that (a) the DOLS estimator is asymptotically efficient if e_t is I(0), (b) the GLS corrected estimator is consistent, but is not as efficient as the DOLS estimator if e_t is I(0), (c) the DOLS estimator is inconsistent if e_t is I(1), and (d) the GLS corrected estimator is consistent if e_t is I(1).

These observations naturally lead to the idea of testing for cointegration by comparing the DOLS estimates and the GLS corrected estimates for β. Let the Hausman-type cointegration be defined by

$$H_T = T \left(\widehat{\beta}_{dgls} - \widehat{\beta}_{dols} \right)' \widehat{V}_{\beta}^{-1} \left(\widehat{\beta}_{dgls} - \widehat{\beta}_{dols} \right) \tag{8}$$

where T stands for the sample size, $\widehat{\beta}_{dols}$ stands for DOLS estimator in level regression, $\widehat{\beta}_{dgls}$ stands for GLS corrected estimator in differenced regression, and $\widehat{V}_{\beta}$ stands for a consistent estimator for the asymptotic variance of $\sqrt{T}(\widehat{\beta}_{dgls} - \beta)$. Under the null hypothesis that the error term is I(0), both estimators $\widehat{\beta}_{dols}$ and $\widehat{\beta}_{dgls}$ are consistent and, therefore, they should be 'close' to each other. The test statistic, H_T, has an asymptotic chi-square distribution with 2 degrees of freedom. On the other hand, under the alternative hypothesis that the error term is I(1), the level regression will be spurious and, therefore, only the differenced regression will be consistent. Therefore, the estimates from these two

[6] We implicitly assume that the error has finite second moments and that the design matrix is of full column rank given the realization of the regressors.

estimators will be very different with a large probability. The test statistic, H_T, diverges in this case.

5 Empirical Results

5.1 Data

Data on real consumption, the real wage rate and the nominal interest rate are required to estimate equation (5). Since the model assumes that money is required to purchase the consumption good, it is more appropriate to use data on those components of consumption that are likely to be purchased by cash, rather than the aggregate consumption expenditure. Cooley and Ogaki (1996) also recommend that at least a component of the aggregate consumption expenditure should be omitted. In this paper, the nondurable- and food-consumption components are used as proxies for the cash good.

We try to select economies with a wide range of inflation experiences for which the relevant consumption data are available. While such consumption data are readily available for developed countries, it is generally not possible to obtain a sufficiently long time-series for most developing countries. Our dataset, comprising a total of 12 countries, is therefore skewed towards the developed economies.[7] The primary sources of data are the *National Accounts of OECD Countries*, the *International Financial Statistics* published by the IMF, and the *United Nations Statistical Yearbook*. Further details regarding the data are provided in the Appendix.

Table 1 presents the summary statistics of the inflation history of these countries over the past two or three decades. The average inflation rate varies from a low of 4.5% for Japan to a high of 90.4% for Israel. The “High” and “Low” columns indicate the variability of the inflation rate as measured by its range.[8] Consistent with the well documented stylized facts in the empirical literature on inflation, higher inflation rates also tend to be associated with a greater variability in the inflation rate. The countries are classified into 3 groups of high- (greater than 10%), medium- (between 5 and 10%) and low- (below 5%) inflation economies to study how the existence and severity of the monetary distortion vary with the inflation rate.

5.2 Trend Properties of the Data

Prior to estimating the cointegrating regressions between real consumption, real wage rates and the gross nominal rates, it is necessary to assess the evidence for the two assumptions that are being made

[7] The 12 countries are Canada, France, Greece, Hong Kong, India, Israel, Italy, Japan, Philippines, Spain, UK and USA.

[8] The range of the inflation rate is defined as the difference between the highest and lowest inflation rates over the sample period.

regarding the trend properties of the data. The first assumption is that all three variables are unit-root nonstationary, which is a pre-requisite for estimating a cointegrating regression. The second assumption is that the two independent variables, the real wage rate and the gross nominal interest rate, are not stochastically cointegrated with each other. If the second assumption is violated, one can still estimate a modified version of Equation (5), but the two parameters of the model can no longer be identified.

Table 2 reports the results of testing the null hypothesis of a unit root, against the alternative of trend-stationarity, based on the Said-Dickey (1984) and the Phillips-Perron (1988) t-ratio tests.[9] At least one of the two tests fails to reject the null of a unit root for most of the variables. Exceptions are the nondurable-consumption for Greece ($\alpha = 10\%$), the Philippine nominal interest rate ($\alpha = 10\%$), Indian food-consumption and real wage rate ($\alpha = 1\%$), and Japanese nondurable-consumption ($\alpha = 5\%$). These results are consistent with Ogaki (1992) and Cooley and Ogaki (1996), who also find evidence in favor of the unit-root hypothesis for food- and nondurable-consumption and the real wage rate.

Table 3 reports the results of the tests for the null hypothesis of no stochastic cointegration between the real wage rate and the gross nominal interest rate. In addition to the Said-Dickey t-ratio test, Park's (1990) $I(1, 5)$ test is also employed. Both tests are based on residuals from an OLS cointegrating regression between the real wage rate and the gross nominal interest rate that includes a time trend. The $I(1, 5)$ test does not reject the null hypothesis for any of the countries at conventional significance levels. However, for three countries (France, India and Japan), the Said-Dickey test is significant ($\alpha = 1\%$) and does not agree with the $I(1, 5)$ test.

Overall, the two assumptions regarding the trend properties of the variables are supported empirically.

5.3 Cointegration Results

Having established that the underlying assumptions are plausible, we proceed to test the empirical validity of equation (5), which embodies the long-run restriction between the stochastic and deterministic trends of real nondurable/food-consumption, the real wage rate and the gross nominal interest rate implied by the model.

Table 4 reports the results of estimating equation (5) using the CCR procedure with nondurable-consumption as the cash good. The first panel reports the results for the group of high-inflation countries. For all four countries, the coefficient of the real wage rate, which measures the intertemporal elasticity of substitution, has the theoretically correct positive sign and is statistically significant at conventional significance levels. The focus of this paper is on the coefficient of the interest rate. It also has the

[9] The Said-Dickey test is also popularly known as the Augmented Dickey-Fuller test.

theoretically predicted negative sign and is statistically significant for all four countries. The point estimates of the interest rate coefficient for Greece and Spain imply that a 1% permanent increase in inflation reduces nondurable-consumption by more than 2% in the long run. The corresponding reduction in nondurable-consumption for the Philippines and Israel is more modest, at 0.9% and 0.2% respectively. With the exception of the $H(1, 3)$ statistic for the Philippines, which is significant at the 1% level, the $H(1, 2)$ and $H(1, 3)$ test statistics do not reject the null hypothesis of stochastic cointegration at conventional significance levels for these 4 countries. The deterministic cointegration restriction is satisfied for all four countries ($\alpha = 1\%$). This is strong evidence in favor of the model.

The second panel of Table 4 reports analogous results for the medium-inflation group of countries. The intertemporal elasticity of substitution has the expected positive sign and is statistically significant for all countries except India, for which it is significantly negative ($\alpha = 5\%$). Possible explanations for the incorrect sign for India are the trend stationarity of the real wage rate and nondurable-consumption, or stochastic cointegration between the real wage rate and the nominal interest rate, which makes the coefficients unidentified. The interest rate coefficient has the predicted negative sign and is statistically significant for France, Hong Kong and Italy ($\alpha = 5\%$). It is negative but insignificant for India, and significant but positive for the UK. The $H(1, 2)$ and $H(1, 3)$ test statistics decisively reject the null hypothesis of no stochastic cointegration for Hong Kong, and are also significant for the UK ($\alpha = 5\%$). They are not significant for France, India and Italy ($\alpha = 5\%$). The deterministic cointegration restriction is strongly rejected for Hong Kong, but not for other countries ($\alpha = 1\%$). Overall, the results for the medium-inflation group are somewhat mixed, with only France and Italy finding clear empirical support.

The last panel of Table 4 reports the results for the group of low-inflation countries. The intertemporal elasticity of substitution is significant and positive only for Japan ($\alpha = 5\%$). The interest rate coefficient has the incorrect positive sign and is statistically significant for all three countries ($\alpha = 1\%$). The $H(1, 2)$ and $H(1, 3)$ statistics are not significant for any of the three countries, and the $H(0, 1)$ statistic is significant only for the US ($\alpha = 1\%$). In contrast to the high- and medium-inflation groups, there is no evidence of monetary distortions for group of low-inflation economies.

Table 5 reports the results of estimating equation (5) using the CCR procedure with food as the cash good. The first panel reports the results for the high-inflation group of countries. The intertemporal elasticity of substitution is correctly signed and statistically significant for all countries except Israel at conventional significance levels. The interest rate coefficient has the expected negative sign and is significant for all 4 countries. The null hypothesis of stochastic cointegration is rejected for Spain ($\alpha = 1\%$) by the $H(1, 2)$ and $H(1, 3)$ statistics, but not for Greece, Israel and the Philippines ($\alpha = 5\%$). The deterministic cointegration restriction is not rejected for any country at the 1% level of significance. These results are similar to those for nondurable-consumption, and support the model's key prediction of monetary distortions.

The second panel of Table 5 reports the results for medium-inflation countries. With the exception of India, the intertemporal elasticity of substitution for all other countries is estimated with the correct positive sign and is also statistically significant ($\alpha = 5\%$). As mentioned earlier, the incorrect sign for India might be caused by trend-stationarity of some of the variables or due to an identification problem. The coefficient of the interest rate has the correct sign for all 5 countries, but is statistically significant only for Hong Kong and Italy ($\alpha = 5\%$). The $H(1,\ 2)$ and $H(1,\ 3)$ statistics do not reject the null hypothesis of stochastic cointegration for any of the countries ($\alpha = 5\%$). The deterministic cointegration restriction is also not rejected by the $H(0,\ 1)$ statistic. Overall, these results support the existence of monetary distortions, but the magnitude of the distortions is smaller and less significant compared to the high-inflation economies.

The last panel of Table 5 reports the results of estimating equation (5) for the low-inflation group of countries. The intertemporal elasticity of substitution has the correct sign and is statistically significant for Canada and Japan ($\alpha = 5\%$). However, it is significantly negative for the U.S. ($\alpha = 5\%$). The interest rate coefficient is incorrectly signed for all 3 countries and is statistically significant ($\alpha = 10\%$). The $H(1,\ 2)$ and $H(1,\ 3)$ statistics reject the null hypothesis of stochastic cointegration for Japan, but not for Canada and the U.S. ($\alpha = 5\%$). The $H(0,\ 1)$ statistic rejects the deterministic cointegration restriction for the U.S., but not for Canada and Japan ($\alpha = 5\%$). These results are similar to those for nondurable-consumption in that the monetary distortions predicted by the model cannot be detected.

Our results with the dynamic OLS procedure are presented in Tables 6 and 7, with nondurable and food consumption as the cash good, respectively. For most of the countries, the results are qualitatively similar to those from the CCR procedure. One notable exception is the result for Greece: the coefficient for the interest rate has the theoretically incorrect positive sign, and it is statistically significant in Tables 6 and 7. There was evidence against the deterministic cointegration for Greece in Tables 4 and 5. It is possible that the dynamic OLS estimator did not work well because of the failure of the deterministic cointegration restriction. Another notable difference is that the evidence for monetary distortions is not statistically significant for either cash good for France and Hong Kong. The Hausman-type test finds some evidence against cointegration only for Canada and United States.

To summarize, for all high-inflation economies, the evidence indicates that statistically significant monetary distortions exist for both of the cash goods for three of the four countries. The results for Greece are consistent with this with the CCR procedure, but inconsistent with the dynamic OLS procedure. Some of our estimates of the elasticity of consumption of the cash goods with respect to the nominal interest rate exceeds 1 in absolute value for these countries, and are likely to translate into economically significant welfare costs. The evidence for monetary distortions for the medium-inflation economies is relatively mixed, with significant distortions evident for France, Italy and Hong Kong for at least one of the cash goods with the CCR procedure, but only for Italy with the dynamic OLS procedure. There is little evidence for the existence of monetary distortions for India and the UK. The empirical results from the CCR procedure provide stronger support for the model for both high- and mdium-inflation countries. This may be because the lag length needed for the parametric correction in the DOLS procedure is constrained in our relatively small samples. In sharp contrast to the medium- and high-inflation economies, virtually no evidence of monetary distortions is apparent for the low-inflation economies with either cash good and econometric procedure.

6 Conclusions

This paper studies the existence and magnitude of monetary distortions in a model of the consumption-leisure choice. Using nondurable- and food-consumption as cash goods, and leisure as the credit good, we document evidence of statistically and economically significant distortions for economies that have experienced double-digit or high single-digit inflation. There appears to exist a threshold level of the rate of inflation, approximately equal to 5%, below which such distortions cannot be observed. A natural extension of this work would be investigating the existence of these monetary distortions with alternative credit goods, such as durable goods. It is hoped that these results enhance our understanding of the welfare costs of predictable inflation.

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Appendix

This Appendix describes the sources of the data in detail. For Hong Kong, India, Israel, Philippines and Spain, data on total real nondurable consumption expenditure and total real food expenditure were obtained from the *United Nations Statistical Yearbook* (*UNSY*). Nondurable consumption expenditure comprised three categories; (a) Food, beverages and tobacco, (b) clothing and footwear, and (c) Medical and health expenses. For the G-6 economies, the relevant consumption data were taken from the OECD's National Income Accounts, whereas for Greece they were taken from *Data Stream.* Aggregate consumption data were converted to per capita terms using population data from the *International Financial Statistics* (*IFS*) for all 12 countries.

It was not possible to obtain the same nominal interest rate series for all countries. For the U.S. and Canada, we used the six month Treasury bill rate; for France and Japan, the lending rate; for UK, the deposit rate; for Hong Kong, the prime rate; for India, Italy, Philippines and Spain the discount rate; and for Israel, the overall cost of unindexed credit. With the exception of Hong Kong and the U.S., the nominal interest rate for all countries was taken from the *IFS.* HK's nominal interest rate series was taken from *Data Stream*, whereas that for the U.S. came from the *Economic Report of the President.*

Where possible, a real wage rate index for the manufacturing sector was used. No real wage data were available for India, Israel and the Philippines. Nominal wage data for these countries was taken from the *UNSY*, and deflated by the *CPI* (from the *IFS*) to yield the real wage rate. The real wage rate index for Hong Kong was taken from the HK government's Census and Statistics Department, whereas that for Spain came from *Data Stream.* Real wage rate indices (for the manufacturing sector) for Canada, France, Italy, Japan and the UK were taken from *Data Stream*, whereas for the U.S. we used the real wage rate for private nonagricultural industries from the *Economic Report of the President.*

Table 1: **Summary Statistics of the Annual Inflation Rate**

Country/ Sample	Average	High	Low
	High-Inflation Countries		
GRC (1961-1995)	11.221	23.826	-0.004
ISR (1979-1994)	90.404	373.820	10.940
PHL (1980-1993)	13.492	46.673	-0.325
SPN (1975-1994)	11.280	24.540	4.570
	Medium-Inflation Countries		
FRA (1973-1998)	6.315	13.749	0.749
HKG (1982-1994)	8.295	11.600	3.440
IND (1975-1991)	7.602	13.870	-7.630
ITL (1970-1997)	9.748	21.277	2.043
UK (1961-1997)	7.217	24.235	1.564
	Low-Inflation Countries		
CAN (1961-1998)	4.820	12.462	0.185
JPN (1970-1997)	4.572	23.122	-0.092
US (1961-1998)	4.634	13.509	1.075

* Countries with average annual inflation rate greater than 10% are classified as "High-Inflation", those with average annual inflation rate between 5% and 10% are classified as "Medium-Inflation", and those with average annual inflation rate less than 5% are classified as "Low-Inflation" economies.

Table 2: **Trend Properties of the Data: Unit Root Tests**

Country	Nondurable		Food		Real Wage Rate		Interest Rate	
	SD/ADF[a]	*PP*[b]	*SD/ADF*[a]	*PP*[b]	*SD/ADF*[a]	*PP*[b]	*SD/ADF*[a]	*PP*[b]
				High-Inflation Countries				
GRC	-3.437*	-3.452*	-1.821	-1.997	-0.817	-0.859	-0.671	0.410
ISR	-1.736	-1.830	-1.861	-2.010	-1.939	-2.057	-2.595	-2.603
PHL	-1.093	-1.460	-2.063	-1.737	-1.875	-1.922	-3.386**	-3.296*
SPN	-2.273	-1.105	-2.173	-1.870	-3.299*	-0.056	-1.153	-0.937
				Medium-Inflation Countries				
FRA	-2.235	-2.294	-2.541	-2.698	-4.358*	-2.205	-0.477	-0.810
HKG	-5.470***	-2.661	-2.312	-2.158	-1.760	-1.436	-2.949	-2.416
IND	-1.100	-4.371**	-5.783***	-6.606***	-9.865***	-12.562***	-2.687	-2.664
ITL	-0.624	-0.584	-0.211	0.240	-0.189	-0.307	-0.535	-0.208
UK	-3.788**	-2.583	-1.424	-1.547	-2.141	-2.225	-1.879	-1.831
				Low-Inflation Countries				
CAN	-1.710	-1.508	-1.772	-1.696	-1.561	-1.049	-1.736	-1.724
JPN	-4.336**	-3.654**	-2.069	-1.932	-3.807**	-5.498***	-2.679	-1.814
US	-2.367	-2.360	-1.655	-1.597	-3.073	-2.748	-2.529	-1.882

[a] *SD/ADF* denotes the Said-Dickey/Augmented Dickey-Fuller t-ratio test for the null hypothesis of a unit root against the alternative of trend-stationarity. The test was performed by starting with three lags and reducing the number of lags until the last lag is significant at the five-percent level. The critical values used incorporate finite sample adjustments based on MacKinnon (1992).

[b] Denotes the Phillips-Perron t-ratio test for the null hypothesis of a unit root against the alternative of trend-stationarity.

[c] *, **, and *** denote significance at the 10%, 5% and 1% levels,respectively.

Table 3: **Trend Properties of the Data: Tests for the Null Hypothesis of No Stochastic Cointegration Between Real Wage Rate and Gross Nominal Interest Rate**

Country	$I(1,5)^a$	SD/ADF^b
	High-Inflation Countries	
GRC	4.254	-3.060
ISR	1.100	-2.553
PHL	3.087	-1.819
SPN	222.420	-2.753
	Medium-Inflation Countries	
FRA	2.821	-7.547***
HKG	12.774	-2.204
IND	8.301	-6.320***
ITL	6.259	-2.970
UK	10.730	-2.915
	Low-Inflation Countries	
CAN	1.171	-2.702
JPN	7.400	-4.874***
US	0.974	-2.349

[a] $I(1,5)$ denotes Park's (1990) test for the null hypothesis of no cointegration. The 1%, 5% and 10% critical values are 0.1027, 0.2506 and 0.4984, respectively.

[b] *SD/ADF* denotes the Said-Dickey/Augmented Dickey-Fuller t-ratio test for the null hypothesis of no stochastic cointegration. The critical values used incorporate finite sample adjustments based on MacKinnon (1992).

[c] *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively.

Table 4: **Canonical Cointegrating Regressions for Real Per Capita Nondurable Consumption**

Country/Sample	$\ln(W_t)^a$	$\ln(1+i_t)^a$	$H(0,1)^b$	$H(1,2)^b$	$H(1,3)^b$
		High-Inflation Countries			
GRC	0.691	-2.063	4.528	1.267	1.393
(1970-1995)	(0.173)	(0.733)	(0.033)	(0.260)	(0.498)
ISR	0.932	-0.191	5.196	0.116	3.063
(1979-1994)	(0.294)	(0.054)	(0.023)	(0.733)	(0.216)
PHL	0.274	-0.925	1.609	2.921	9.845
(1980-1993)	(0.055)	(0.300)	(0.205)	(0.087)	(0.007)
SPN	0.173	-2.331	1.735	0.129	1.469
(1975-1994)	(0.032)	(0.511)	(0.188)	(0.719)	(0.480)
		Medium-Inflation Countries			
FRA	0.962	-1.257	0.507	0.038	4.806
(1973-1998)	(0.055)	(0.252)	(0.477)	(0.846)	(0.090)
HKG	2.473	-0.738	33.625	13.053	151.734
(1982-1994)	(0.068)	(0.113)	(0.000)	(0.000)	(0.000)
IND	-0.518	-0.615	1.469	2.424	3.373
(1975-1991)	(0.048)	(0.424)	(0.225)	(0.119)	(0.185)
ITL	0.489	-0.477	0.258	3.449	3.806
(1970-1997)	(0.051)	(0.233)	(0.611)	(0.063)	(0.149)
UK	0.173	1.139	0.001	4.553	8.677
(1961-1997)	(0.035)	(0.241)	(0.977)	(0.033)	(0.013)
		Low-Inflation Countries			
CAN	0.067	1.988	0.184	2.143	3.769
(1961-1998)	(0.088)	(0.470)	(0.668)	(0.143)	(0.152)
JPN	0.410	1.715	1.708	0.721	2.208
(1970-1997)	(0.024)	(0.249)	(0.191)	(0.396)	(0.332)
US	-0.048	1.847	6.880	0.000	0.179
(1961-1998)	(0.474)	(0.924)	(0.009)	(0.993)	(0.914)

[a] Standard errors are in parenthesis.

[b] H(0,1) tests the deterministic cointegration restriction, whereas H(1, 2) and H(1,3) test the null hypothesis of stochastic cointegration. P-values are in parenthesis.

Table 5: **Canonical Cointegrating Regressions for Real Per Capita Food Consumption**

Country/Sample	$\ln(W_t)^a$	$\ln(1+i_t)^a$	$H(0,1)^b$	$H(1,2)^b$	$H(1,3)^b$
		High-Inflation Countries			
GRC	0.890	-2.659	10.832	1.828	3.276
(1962-1995)	(0.209)	(1.031)	(0.001)	(0.176)	(0.348)
ISR	0.137	-0.143	3.062	0.155	4.090
(1979-1994)	(0.265)	(0.053)	(0.080)	(0.694)	(0.130)
PHL	0.210	-0.347	1.199	2.659	5.241
(1980-1993)	(0.046)	(0.270)	(0.274)	(0.103)	(0.073)
SPN	0.090	-0.960	0.000	9.534	9.645
(1975-1994)	(0.015)	(0.216)	(0.999)	(0.002)	(0.008)
		Medium-Inflation Countries			
FRA	0.576	-0.117	0.295	1.843	5.285
(1973-1998)	(0.054)	(0.231)	(0.587)	(0.175)	(0.071)
HKG	0.589	-0.216	0.178	0.253	5.912
(1982-1994)	(0.019)	(0.042)	(0.673)	(0.615)	(0.052)
IND	-0.344	-0.243	0.697	2.216	3.429
(1975-1991)	(0.047)	(0.393)	(0.406)	(0.136)	(0.180)
ITL	0.833	-1.312	0.045	0.275	1.898
(1970-1997)	(0.056)	(0.258)	(0.832)	(0.600)	(0.387)
UK	0.802	-0.063	2.821	3.260	3.548
(1961-1997)	(0.019)	(0.106)	(0.093)	(0.071)	(0.170)
		Low-Inflation Countries			
CAN	0.658	0.376	0.271	1.419	1.566
(1961-1998)	(0.036)	(0.110)	(0.603)	(0.233)	(0.457)
JPN	0.698	0.735	2.217	5.044	15.441
(1970-1997)	(0.024)	(0.252)	(0.136)	(0.025)	(0.000)
US	-2.060	3.326	7.343	0.083	3.097
(1961-1998)	(0.941)	(1.756)	(0.007)	(0.773)	(0.213)

[a] Standard errors are in parenthesis.

[b] H(0,1) tests the deterministic cointegration restriction, whereas H(1, 2) and H(1,3) test the null hypothesis of stochastic cointegration. P-values are in parenthesis.

Table 6: **Hausman-type Cointegration Test for Real Per Capita Nondurable Consumption**

Country	Lead & Lag[a]	DOLS[b]		Hausman
		$\ln(W_t)$	$\ln(1+i_t)$	Test
		High-Inflation Countries		
GRC	2	-0.0063	1.8005	1.4512
(1962-1995)		(0.1118)	(0.6710)	
ISR	1	0.4718	-0.1090	0.5061
(1979-1994)		(0.5576)	(0.0574)	
PHL	1	0.3871	-1.0402	0.1995
(1980-1993)		(0.0481)	(0.3653)	
SPN	2	0.3145	-2.2321	0.2951
(1975-1994)		(0.0179)	(0.1488)	
		Medium-Inflation Countries		
FRA	2	1.2862	-0.6140	1.2244
(1973-1998)		(0.7642)	(1.7822)	
HKG	0	2.2564	-0.7280	1.3848
(1982-1994)		(0.2371)	(0.9424)	
IND	1	-0.3427	6.9375	0.0287
(1975-1991)		(0.0516)	(1.3000)	
ITL	2	0.5605	-1.0582	2.0123
(1970-1997)		(0.2011)	(0.2953)	
UK	2	0.7891	0.1339	0.2154
(1961-1997)		(0.0281)	(0.1754)	
		Low-Inflation Countries		
CAN	2	0.2610	1.0502	6.0997**
(1961-1998)		(0.1144)	(0.3839)	
JPN	2	0.7087	0.5126	0.3853
(1970-1997)		(0.2514)	(2.4783)	
US	2	-0.4206	-1.1938	16.3383***
(1961-1998)		(0.5445)	(1.5429)	

[a] Lead and lag lengths of Hausman-type Cointegration test is selected by the BIC method with maximum length is set to 2.

[b] Standard errors, calculated by the Newey-West HAC estimator with bandwidth parameter of 2, are in parenthesis.

[c] Critical value of $\chi^2(2)$ are 4.61, 5.99 and 9.21 for 10%, 5% and 1% significance levels. *, ** and *** denote the rejection at 10%, 5% and 1% significance levels.

Table 7: **Hausman-type Cointegration Test for Real Per Capita Food Consumption**

Country	Lead & Lag[a]	DOLS[b]		Hausman
		$\ln(W_t)$	$\ln(1+i_t)$	Test
		High-Inflation Countries		
GRC	2	0.0567	2.3733	0.9985
(1962-1995)		(0.0489)	(0.2933)	
ISR	1	0.6355	-0.0540	0.4511
(1979-1994)		(0.3388)	(0.0348)	
PHL	1	0.3712	-0.9749	0.1935
(1980-1993)		(0.0502)	(0.3807)	
SPN	2	0.2427	-0.1609	0.3105
(1975-1994)		(0.0087)	(0.0720)	
		Medium-Inflation Countries		
FRA	2	0.3397	-0.8032	0.5177
(1973-1998)		(0.6831)	(1.5929)	
HKG	0	0.6198	-0.2821	0.3560
(1982-1994)		(0.0619)	(0.2461)	
IND	1	-0.2687	6.3242	0.0409
(1975-1991)		(0.0634)	(1.5970)	
ITL	2	0.2088	-0.1522	1.4627
(1970-1997)		(0.1381)	(0.2028)	
UK	2	0.1555	1.0196	1.2702
(1961-1997)		(0.0600)	(0.3747)	
		Low-Inflation Countries		
CAN	2	-0.2875	0.8585	5.7401*
(1961-1998)		(0.1340)	(0.4497)	
JPN	2	0.3901	0.9048	0.3158
(1970-1997)		(0.1050)	(1.0356)	
US	2	-0.0756	-0.8063	8.4532**
(1961-1998)		(0.3527)	(0.9996)	

[a] Lead and lag lengths of Hausman-type Cointegration test is selected by the BIC method with maximum length is set to 2.

[b] Standard errors, calculated by the Newey-West HAC estimator with bandwidth parameter of 2, are in parenthesis.

[c] Critical value of $\chi^2(2)$ are 4.61, 5.99 and 9.21 for 10%, 5% and 1% significance levels. *, ** and *** denote the rejection at 10%, 5% and 1% significance levels.