A Microeconometric Comparison of Household Behavior Between Countries

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This article reviews the methodological issues associated with estimating and testing competitive markets with panel data and applies it to an important new arena, international economics. We investigate differences in housing consumption and male labor supply across households in Germany and the United States with data from the German Socio Economic Panel and the Panel Study of Income Dynamics. Observations are treated as equilibrium outcomes, generated by a set of competitive markets that accommodate varying degrees of integration. After controlling for differences in the sample populations and aggregate shocks, we find that the estimated structural parameters are surprisingly similar across these two countries.

KEY WORDS: Complete markets; Generalized method of moments; International comparisons; Intertemporal labor supply; Panel data; PPP.

In recent years the integration of financial, commodity, and labor markets has drastically increased. This integration can be attributed to a variety of factors, such as reduced transportation costs, innovations in the communication and computer sectors, the emergence of transnational corporations that have reduced technology gaps between countries, the deregulation of financial markets, and the steady dismantling of trade and foreign investment barriers, undertaken both unilaterally and through agreements. In part as a result of these developments, the United States and other industrialized countries have undergone major changes in their wage structure and income distribution (e.g., see Mincer 1991; Davis 1992; Katz and Murphy 1992).

This article studies the degree to which the integration of markets is reflected at the household level. Broadly speaking, trade and factor price equalization theorems have established that, if there is enough trade in consumption commodities and/or factors in production, the prices of those goods and services that are not traded across national boundaries will also be equalized globally. A powerful test of integration, then, is to focus on goods and services that are not normally traded on international markets (say because of regulation or natural barriers). To this end we conduct an empirical study of the male leisure and housing consumption by married couples in the United States and Germany with panel data. These services are typically most substantive in a typical household's budget, especially if one follows Becker (1971) in taking a full-income approach to household income. This might be one reason why housing consumption and male labor supply are two of only a very few goods and services collected by both the Panel Study of Income Dynamics (PSID), and the German Socio Economic Panel (GSOEP), on which our study is based.

We organize the data around the null hypothesis that competitive and complete markets (CCM) exist, an assumption first tested by Altug and Miller (1990). Their framework, which is a natural generalization of earlier work on perfect foresight models of life-cycle behavior, yields a tractable econometric model with a simple factor structure. [Heckman and MaCurdy (1980), MaCurdy (1981, 1983), Altonji (1986), and Browning, Deaton, and Irish (1985) are representative of authors of econometric models from this earlier literature on life-cycle behavior.] We shall demonstrate that the CCM assumption is notoriously difficult to relax, without losing the coherent economic interpretation that accompanies it. Although Altug and Miller (1990) only found limited evidence against CCM, several studies later have rejected this hypothesis. Thus, by way of critical review, we reexamine tests of CCM, as well as a closely related literature on whether consumption is more volatile than the permanent income hypothesis (PIH) would predict. (See Hall and Mishkin 1982; Altonji and Siow 1987; Zeldes 1989; Runkle 1991; Mariger and Shaw 1993; Nelson 1994; Lusardi 1996.) Therefore our empirical analysis can be interpreted as a further investigation of CCM, bringing new data to bear on that hypothesis.

The second motivation for our work is to compare the two economies of Germany and the United States over the last several years at the micro level. Macro data exhibit differences in labor supplied, housing demanded, and wages and rents paid across the two countries. The question we pose is whether these differences are significant and to what extent they can be explained by country-specific differences in preferences and technology, differences in the composition of populations, or differences in the price structures for services and commodities.

The article is organized as follows. Section 1 lays out the general equilibrium model that provides the basis for our approach. At the same time we review the growing literature on the use of dynamic structural models to explain the equilibrium relationship between consumption and leisure allocations observed in cross-sectional and panel datasets. Section 2 introduces the parameterization of our model and presents our estimation strategy. The data were taken from

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the PSID and the GSOEP and are described in Section 3. Our empirical findings are presented and discussed in Section 4. A brief conclusion draws together our analysis.

1. FRAMEWORK

1.1 The Permanent Income Hypothesis

Consider the following (global) economy. Production and consumption occur on discrete dates $t \in \{1, 2, \ldots\}$. Let Ω denote the probability space that characterizes the possible paths or histories the economy might take, and let $\omega \in \Omega$ denote a generic outcome. Information about the economy is modeled as a sequence of σ algebras, denoted by $\ldots F_t \subseteq F_{t+1} \ldots$, where F_t is interpreted as the information available at time t. Commodities consumed by household n at time t are denoted by $c_{nt}(\omega)$. Conditional on history ω pertaining, choices at t can only be based on what is known at that time. This implies that they must be measurable with respect to F_t. For notational reasons we abbreviate $c_{nt}(\omega)$ by c_{nt}. Similarly, labor supply by household n at time t is an F_t measurable function, which is denoted by $l_{nt}(\omega)$, or l_{nt} for short. Except for the labor market, the production side of the economy is ignored. To incorporate discounting over time, we allow the within-period utility function, $U_t(\cdot)$, to be a time-dependent function. Given any value of (l_{nt}, c_{nt}), we also assume that $U_t(\cdot)$ is F_t measurable to reflect exogenous stochastic demographic factors. We assume that household n obeys the expected utility hypothesis, maximizing the utility functional

$$E_t\left\{\sum_{t=t_1}^{t_2} U_t(l_{nt}, c_{nt})\right\} \tag{1}$$

by trading its endowments, including labor, at given prices. (Thus, we assume that markets are competitive, at least so far as consumers and workers are concerned, ignoring, for example, the monopolistic behavior by unions.) Accordingly, let $q_{rnt} = q_{rnt}(\omega)$ denote the quantity of asset r held by household n in period t if history ω occurs, $p_{rt} = p_{rt}(\omega)$ its price, and $s_{rt} = s_{rt}(\omega)$ the associated dividend. In addition to constraining bequests (to be nonnegative, for example), we require, that for each date t and history $\omega \in \Omega$,

$$\sum_{r\in R} [p_{rt}(q_{rnt} - q_{rnt+1}) + q_{rnt}s_{rt}] \geq c_{nt} - w_{nt}l_{nt}, \tag{2}$$

where $w_{nt} = w_{nt}(\omega)$ denotes the wage rate.

This framework, which is often referred to as the permanent income hypothesis (PIH), yields two kinds of first-order conditions. Because there is a spot market for labor services, agents equate their marginal rate of substitution between leisure and goods to the real wage rate (assuming an interior solution):

$$\frac{\partial U_t(c_{nt}, l_{nt})/\partial l_{nt}}{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}} = w_{nt}. \tag{3}$$

This condition relates the pairwise allocation of currently traded commodities. The other first-order condition for this problem equates the relative price of a good to be received next period in several states of the world with its matching marginal rate of substitution. Letting $\pi_{rt} = \pi_{rt}(\omega) = [p_{rt+1}(\omega) + s_{rt+1}(\omega)]/p_{rt}(\omega)$ denote the return on asset r in time t, the marginal rate of substitution between current consumption and future consumption is

$$E_t\left\{\pi_{rt}\,\frac{\partial U_{t+1}(c_{nt+1}, l_{nt+1})/\partial c_{nt+1}}{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}}\right\} = 1. \tag{4}$$

To interpret (4), suppose, for example, that π_{rt} paid off two units in good states of the world, one unit in medium states, but nothing in bad states. Then (4) would be an equilibrium condition for the marginal rate of substitution between a unit of today's consumption and an uncertain amount to be delivered tomorrow (2 if the good states occur, 1 if the medium states occur, and 0 otherwise) with the price π_{rt}. Unless enough assets exist (can be constructed) so that each one pays off in only one state of the world, consumers cannot choose an optimal consumption bundle that equates their marginal rate of substitution in a pairwise fashion across the consumption space. Limited market opportunities compel the consumer to buy bundles of commodities that are typically intertemporally linked through his/her particular asset portfolio, even if his/her own utility function is additively separable across time. As we shall see, this wreaks havoc on the identification and estimation of preferences. Altug and Miller (1990), Altug and Labadie (1994), and Card (1994, pp. 49–78) recognized this problem, and readers are referred to their respective discussions. Our approach, however, is complementary to theirs, differing in the way we cast the identification problem under PIH.

Several studies attempt to test the PIH with panel data based on Equation (4). In the spirit of this literature, let ε_{rnt} denote the difference between the expectation on the left side of (4) and its realization:

$$\varepsilon_{rnt} = E_t\left\{\pi_{rt}\,\frac{\partial U_{t+1}(c_{nt+1}, l_{nt+1})/\partial c_{nt+1}}{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}}\right\} - \pi_{rt}\,\frac{\partial U_{t+1}(c_{nt+1}, l_{nt+1})/\partial c_{nt+1}}{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}}, \tag{5}$$

where $E_t[\varepsilon_{rnt}] = 0$. Substituting (5) into (4), we obtain, in logarithmic form, the following identity for all (r, n, t) and ω:

$$\Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}] = \ln(1 - \varepsilon_{rnt}) - \ln(\pi_{rt}), \tag{6}$$

assuming that $|\varepsilon_{rnt}| < 1$ and writing Δ for the difference operator between an object's value today and the value it takes the following period. Suppose that we now average (6) over the cross-section and define $v_{rt}^{(0)}$ as the limit in the sample size N. (Strictly speaking we should provide conditions that guarantee that this limit exists. A sufficient condition is that the data are a random sample of the population.) That is,

$$\begin{aligned} v_{rt}^{(0)} &= \text{plim}_{N\to\infty}\frac{1}{N}\sum_{n=1}^{N}[\ln(1-\varepsilon_{rnt}) - \ln(\pi_{rt})] \\ &= \text{plim}_{N\to\infty}\frac{1}{N}\sum_{n=1}^{N}\Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}]. \end{aligned} \tag{7}$$

The estimation approach is to treat $v_{rt}^{(0)}$ as a time dummy to be estimated in the second line of (7) and form an orthogonality condition in a cross-section or panel context. Similarly, one could multiply Equation (6) by an instrument vector $z_{nt} = (z_{nt}^1, \ldots, z_{nt}^q)'$, define a vector of dummy variables v_{rt} as

$$v_{rt} = \operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} z_{nt}[\ln(1-\varepsilon_{rnt}) - \ln(\pi_{rt})], \quad (8)$$

and obtain q orthogonality conditions of the form

$$\operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} z_{nt} \Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}] - v_{rt} = 0. \quad (9)$$

Because the q-dimensional coefficient on the time-dummy vector must be estimated if (9) is to be exploited in estimation, this is of no help in identifying the structural parameters of interest unless they are restricted in some way. Without such restrictions, there are more parameters to estimate than orthogonality conditions. As a practical matter, the studies essentially assume that

$$v_{rt} = v_{rt}^{(0)} \operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} z_{nt}. \quad (10)$$

Substitute (10) into (9) and minimize a quadratic criterion function in

$$\frac{1}{N} \sum_{n=1}^{N} z_{nt}\{\Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}] - v_{rt}^{(0)}\} \quad (11)$$

by choosing one time-dummy coefficient per year and the structural parameters of interest. From (8), Assumption (10) is equivalent to assuming that

$$\operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} z_{nt} \ln(1-\varepsilon_{rnt}) = \operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} z_{nt} \times \operatorname{plim}_{N\to\infty} \frac{1}{N} \sum_{n=1}^{N} \ln(1-\varepsilon_{rnt}). \quad (12)$$

From a formal modeling perspective, there is no reason to believe that the time dummies obey the restrictions (10), or equivalently that there exists a set of instruments satisfying (12), uncorrelated with the forecast error made by a cross-section of the population at any one point in time. Incomplete markets force agents to make choices over bundles or groups of goods that are not necessarily in the proportions they would desire if the commodities were available separately. Thus, one would expect past allocations and the factors that affected past choices to affect current and future goods allocations as well. The very fact that consumers cannot divorce their choices over current quantities from what is available to them in the future even after controlling for their endowments now and their planned bequests is the reason picking instruments is so difficult for econometricians. It does not seem possible to identify and estimate parameters characterizing the utility function without imposing more assumptions than the PIH does. Intuitively, this should not be all that surprising. A modern statement of the PIH is that households maximize their expected lifetime utility subject to resource constraints. It is a statement about rationality, made irrespective of markets, and, as is well known, equilibrium allocations are very difficult to characterize when markets are incomplete. Apparently recognizing the difficulties associated with finding such instruments, Zeldes (1989) and Runkle (1991) took a second-order expansion around the error term. A straightforward extension of our critique, however, shows that there is no reason to believe that the residual of a second-order approximation is uncorrelated with time-varying individual-specific instruments. Therefore, their approach does not avoid the difficulties described previously. Yet, although nobody would estimate a model of preferences that are nonadditively separable across time without specifying its functional form, researchers seem to ignore the dual problem induced by a nonlinear budget constraint.

1.2 Competitive and Complete Markets

Imposing assumptions on the trading environment can break this identification problem. Specifically, the assumption of complete and competitive markets (CCM) practically guarantees the existence of such instruments. To see this, we follow the standard practice in general equilibrium, established by Debreu (1959) and others, of assuming that all trades take place at date 0 and that the history only determines which path of deliveries and consumptions gets realized. Accordingly, let $\Lambda_t(\omega)$ denote the price measure for commodities consumed in period t when the history is ω, and denote by $\lambda_t = \lambda_t(\omega)$ its Radon Nikodym derivative. That is, the cost of a unit of the consumption vector paid at date 0 to be delivered in the event of $A \in F_t$ at date t is

$$\int_{A_t} \lambda_t(\omega)\, dP. \quad (13)$$

Note that under CCM the household must only obey a lifetime budget constraint of the form

$$E_0 \left\{ \sum_{t=t_1}^{t_2} \lambda_t [c_{nt} - l_{nt} w_{nt}] \right\} \le W_n, \quad (14)$$

where W_n is the present value of wealth minus bequest. This is the weakest constraint the household might face, which merely ensures that total receipts from labor income exceed expenditure outlays over the household lifetime. It implies that the aggregate shocks are fully transmitted through prices of contingent claims. Consequently, as in (3) but in contrast to (4), pairwise comparisons between commodities would be independent of all other commodities but for nonseparabilities in the utility function. First-differencing the logarithm of an interior first-order condition, one obtains

$$\Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}] = \ln(\lambda_{t+1}) - \ln(\lambda_t). \quad (15)$$

The condition states that, in equilibrium, the (logarithm of the) marginal rate of substitution between consumption in two consecutive periods equals (the logarithm of) their relative prices.

Notice that (15) is a much stronger condition than (6), which contains the troublesome forecast error. Admittedly, the PIH equation (6) and the CCM equation (15) have a similar appearance, especially, if one allows for the possibility that there are unobservable factors and measurement error in the variables that make up the marginal rate of substitution function. But as we have shown, basing estimation on (6) is problematic, whereas orthogonality conditions are easy to construct from (15). We offer several interpretive remarks about tests based on (15). First, the aggregation results of Rubinstein (1981) and others show that in addition to actuarially fair insurance against purely idiosyncratic shocks only a few markets are required to exhaust trading opportunities. In a deterministic world, a one-period interest rate plus spot markets suffice. For the utility parameterizations most researchers focus on (the hyperbolic absolute risk aversion class), one additional security suffices. Thus, CCM is not as strong as it looks, given the parametric assumptions that researchers are already making in other parts of their framework. Second, the tests are primarily about the existence of spot markets and resource allocations; lacking data on food prices, for example, these tests do not distinguish between an allocation mechanism due to markets versus an efficiently operated food-stamp program. All that can be identified is whether the status quo equates marginal rates of substitution across people.

Most of the studies following Altug and Miller (1990) claim to reject what is sometimes referred to as full insurance. (See Mace 1991; Cochrane 1991; Townsend 1994; Udry 1994; Ham and Jacobs 1994; Hayashi, Altonji, and Kotlikoff 1996.) There are, of course, many other reasons orthogonality conditions based on (15) might be rejected. None of these tests maintain the null hypothesis of CCM in a vacuum, but they simultaneously make assumptions about the separability of commodities, measurement error, the distribution of the unobservables, stationarity of the series, and, of course, the parameterization itself. Thus it is impossible to reject the hypothesis of CCM; it is only possible to reject all of the maintained hypotheses at once or alternatively characterize the assumptions under which we fail to reject CCM. Indeed testing CCM can be undertaken in a perfect foresight economy; that is, $\omega \in F_0$ and $F_0 = F_1 = \dots$. In this case, consumers simply maximize

$$\sum_{t=t_1}^{t_2} U_t(l_{nt}, c_{nt}) \tag{16}$$

subject to

$$\sum_{t=t_1}^{t_2} \lambda_t(c_{nt} - l_{nt}w_{nt}) \leq W_n, \tag{17}$$

and exactly the same set of first-order conditions and estimation equations emerge. In this case, $(\lambda_t(\omega))^{-1} - 1$ is the interest rate in period t. Undue emphasis on the insurance aspects disguises other valid reasons the framework is misspecified (such as nonseparable preferences). It also presupposes that uncertainty is a central issue in consumption smoothing over time (as opposed to savings and loans, for example).

1.3 Estimating and Testing CCM

To exploit (3) and (15), the empirical papers on CCM have followed the literature on life-cycle consumption by postulating an observed and an unobserved part of the marginal rate of substitution function. Let $E[\Delta \ln(\partial U_t(c_{nt}, l_{nt})/\partial c_{nt})|x_{nt}]$ denote the expected value of the logarithm of the marginal rate of substitution between consumption in successive periods conditional only on an exogenous vector of household characteristics x_{nt}. Then,

$$\Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}] = E[\Delta \ln\{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}\}|x_{nt}] + \varepsilon_{nt}, \tag{18}$$

where, by construction,

$$E[\varepsilon_{nt}|x_{nt}] = 0. \tag{19}$$

By specifying $E[\Delta \ln(\partial U_t(c_{nt}, l_{nt})/\partial c_{nt})|x_{nt}]$ as a known function of observed variables, (l_{nt}, c_{nt}, x_{nt}), and a parameter vector, $\theta_0 \in \Theta$, to be estimated, instrumental variables, z_{nt}, can be chosen that satisfy the condition

$$\operatorname{plim}_{N\to\infty} \frac{1}{N}\sum_{n=1}^{N} z_{nt} \times [E[\Delta \ln\{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}\}|x_{nt}] - \Delta \ln(\lambda_t)] = 0. \tag{20}$$

The main requirement is that the instruments (such as the exogenous household characteristics) are uncorrelated with the unobserved variables in current utility. A similar approach is taken to estimate (3).

In implementing the finite-sample analog to (20), Altug and Miller (1990) exploited (19), which, by the law of large numbers and random sampling, implies the qT orthogonality conditions

$$\frac{1}{N}\sum_{n=1}^{N}\left\{\begin{array}{c} z_{n1}\varepsilon_{n1} \\ \cdots \\ \cdots \\ z_{nT}\varepsilon_{nT} \end{array}\right\}. \tag{21}$$

If there are more sample moments than parameters to estimate, the overidentifying restrictions may be tested, thus providing a specification test of the overall framework.

Their approach parts company with several studies, written both before and afterward (e.g., see Zeldes 1989; Runkle 1991; Ham and Jacobs 1994; Hayashi et al. 1996). These other studies treat each (n, t) observation as an element in the sequence that is summed over T, the number of time periods in the panel, and N, the number of households. The large-sample properties of the resulting estimator depend on the convergence of

$$\frac{1}{TN}\sum_{n=1}^{N}\sum_{t=1}^{T} z_{nt}\varepsilon_{nt}. \tag{22}$$

From (22), it is easy to see that, if the observations are grouped by households, then we can view the summation as an average of $T^{-1}(z_{n1}\varepsilon_{n1} + \cdots + z_{nT}\varepsilon_{nT})$ over n; it immediately follows from the law of large numbers that (22) converges almost surely to 0 and, scaled up by $N^{1/2}$, is asymptotically distributed as a normal random variable. Consider, on the other hand, treating each $z_{nt}\varepsilon_{nt}$ as a separate observation in the summand. Without serial correlation and heteroscedasticity over time—that is, assuming that $E[\varepsilon_{n,t}\varepsilon_{n,t-s}] = 0$ and $E[\varepsilon_{n,t}^2] = E[\varepsilon_{n,t-s}^2]$ for all t and $s \neq 0$—sampling across different time periods has no special significance. Hence, the two approaches generate essentially the same test statistics. [The test statistics are numerically equal if the covariance and variance restrictions noted in the text are imposed when implementing the approach based on (20).] The CCM hypothesis does not require these restrictions to hold, however, and as an empirical matter our study finds these assumptions unjustified. Appendix A exhibits the correlation matrix for the German subsample used in our study. We find substantial serial correlation within a household and over time. The American subsample is not reported but exhibits similar patterns. When these assumptions are unjustified, it follows that the procedure of ignoring the heteroscedasticity and serial correlation will yield biased test statistics and incorrect standard errors.

A second way to test CCM involves adding variables that under the null hypothesis of CCM have no systematic effect. The alternative hypothesis is usually characterized by the statement that consumption is excessively volatile. To implement the test, let y_{nt} be a vector of observed variables that do not enter the observed part of the marginal rate of substitution function, and suppose that we form orthogonality conditions from

$$E[\Delta \ln\{\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}\}|x_{nt}] - \Delta \ln(\lambda_t) + \alpha y_{nt} = 0. \quad (23)$$

Assuming that we can find variables y_{nt} that satisfy the condition

$$\operatorname{plim}_{N\to\infty} \frac{1}{N}\sum_{n=1}^{N} z_{nt}y_{nt} = 0, \quad (24)$$

it follows that the null hypothesis of CCM would be rejected if the estimate of α differs significantly from 0. This approach is informative but should be seen in context. Supposing that y_{nt} is correlated with unobserved variables in the utility function, the estimated coefficient would be nonzero even if the parameterization is correctly specified and CCM holds. In practice, the kinds of variables researchers have looked at are most susceptible to this criticism. For example, if current income is correlated to health, marginal productivity at home might rise as well. If not properly accounted for within the marginal utility of consumption, the estimate of α could be spuriously positive: People who fall ill typically spend less resources on vacations and reduce their work effort too. Another example—because food preparation takes time, a person might be indifferent between working more hours and eating out versus working less hours and preparing meals at home. This also induces a positive correlation between hours at work and unobservables in the marginal utility of food. Tests that ignore such correlations run the risk of falsely rejecting the null hypothesis.

Apart from ignoring potential correlations between marginal utility and income, several studies in this literature are misleading in another respect. They mistakenly interpret the null hypothesis as the PIH rather than CCM, which leads them into a spurious discussion of whether unanticipated increases in income give rise to excessive consumption. Moreover, to calculate the difference between unanticipated and anticipated income, they only condition on a few variables observed by the econometrician rather than the much larger information set of the agent. As we have pointed out, nothing in a CCM framework with time-separable preferences differentiates it from a perfect foresight model. Therefore, unanticipated changes in income cannot be distinguished from anticipated changes using data on income and consumption streams alone.

Recently the life-cycle approach has been criticized for not providing a useful framework for understanding the main features of individual labor supply. In particular, Abowd and Card (1989) argued that the covariance structure between labor supply and earnings can be explained by a simple factor model that includes one permanent factor (which they call a productivity shock) and two other factors reflecting transitory shocks. One property of their factor model is that productivity shocks have the same impact on labor supply and earnings. Their critique imposes strong separability assumptions for within-period consumption. Accordingly, suppose that current utility takes the form

$$U_t(l_{nt}, c_{nt}) = \beta^t[\alpha_0^{-1}\delta_{0t}c_{nt}^{\alpha_0} + \alpha_1^{-1}\delta_{1t}l_{nt}^{\alpha_1}], \quad (25)$$

where $\delta_{it} = \exp(x'_{nt}B_i + \varepsilon_{int})$ for $i \in \{0, 1\}$. Substituting (25) into Equation (15) and solving for the changes in labor supply yields

$$\Delta \ln(l_{nt}) = \frac{1}{(\alpha_1 - 1)}[\ln(\beta) + \Delta\ln(\lambda_t) + \Delta\ln(w_{nt}) - \Delta x'_{nt}B_1 - \Delta\varepsilon_{1nt}]. \quad (26)$$

Changes in the labor income can then be expressed as

$$\begin{aligned}\Delta \ln(w_{nt}l_{nt}) &= \Delta\ln(w_{nt}) + \Delta\ln(l_{nt}) \\ &= \frac{\alpha_1}{(\alpha_1 - 1)}\Delta\ln(w_{nt}) + \frac{1}{(\alpha_1 - 1)} \\ &\quad \times [\ln(\beta) + \Delta\ln(\lambda_t) - \Delta x'_{nt}B_1 - \Delta\varepsilon_{1nt}]. \end{aligned} \quad (27)$$

From Equations (25)–(27), one can see that even a simple life-cycle model has at least two factors, which capture the impact of productivity shocks, if $\alpha_1 < 1$. Relaxing the strong separability assumptions and including female leisure into the current utility function adds another factor. Incorporating tax changes into the framework would add a further shock (see Sieg 1995). Hence, a factor model that includes only one permanent productivity shock is incompatible with the life-cycle model and more generally CCM. Moreover, as α_1 approaches 1, the model reduces to a one-factor model in which hours are highly elastic to wages. But this contradicts the low point estimates typically found by researchers directly estimating the model. The extent to which the tests of Abowd and Card (1989) are powerful against the richer covariance structures that are generated by CCM suggests that life-cycle models are overparameterized. Our study follows Altug and Miller (1990), however, in empirically rejecting a one-factor parameterization.

A direct consequence of the one-factor structure that Abowd and Card (1989) imposed is that, aside from idiosyncratic transitory components, labor supply and earnings roughly move together. Therefore, most variation in hours cannot be attributed to wage changes. There are two possible explanations for this stylized fact. First, factors that directly influence labor supply but not wages will yield this proportionality result provided that systematic wage changes are small compared to overall changes in these factors throughout the sample phase. Such factors include the price for contingent claims, λ_t, aggregate shifts observed in the data of the exogenous characteristics of the household, x'_{nt}, and aggregate shifts in female labor supply, l_{0nt} (i.e., if it enters the within-period utility function in a nonseparable way). Second, markets may be incomplete so that productivity changes affect hours but not paid wages. The thrust of our argument is that many tests of CCM do not incorporate a sufficiently rich specification of the model and hence do not adequately control for heterogeneity throughout the population.

1.4 Relaxing the CCM Null Hypothesis

The upshot of this discussion is that CCM does indeed provide a useful framework for thinking about data. But, as with any null, the failure to reject it may merely be poor data. A rejection, on the other hand, should be interpreted as a strike against all the assumptions that bolster the framework, including those that are entirely unrelated to the questions of market structure. One way to overcome the limitations of this approach is to relax CCM by explicitly modeling the rationale for incomplete markets or by relaxing the strong separability assumptions about preferences that underpin (15). For example, in an empirical study of managerial compensation, Margiotta and Miller (1994) explicitly modeled the additional constraints imposed on shareholders when actions taken by their firm's manager is private information. Card (1990) estimated a model with a minimum-hours threshhold (discussed by Miller 1990, pp. 169–180). Altug and Miller (1994) derived a set of equilibrium conditions that identify the structural parameters of interest when preferences are not separable over time. Sieg (1995) included income taxation into the model and analyzed how changes in the tax structure affect household choices and the distribution of the tax burden.

Tests of CCM can be specialized to particular markets, however, or to specific populations. From (15) it is clear that the only markets on which data exist are being tested; consequently, this test has no power against goods and services that are not observed. Another way of estimating a more limited set of complete markets is to assume that the gains from trade are only exhausted within a subpopulation rather than over the whole. Hayashi et al. (1996) and Townsend (1994) investigated the restrictions implied by CCM within dynastic families and Indian villages, respectively. Hayashi et al. (1995) used the framework of Altug and Miller (1990) to conduct their empirical analysis. Townsend (1994) imposed three additional restrictions. First, he did not exploit data on time-varying characteristics within the population. Second, he imposed additive separability between goods and services consumed concurrently. Third, his choice of the utility function from the constant absolute risk-aversion class implies that the Frisch demand has a closed-form solution. Imposing these three restrictions enabled Townsend to express λ_t as a mapping of aggregate consumption. Although Townsend did not test them himself, the first two requirements of his approach are typically violated in panel datasets (see Browning and Meghir 1991; Altug and Labadie 1994).

To test these more limited versions of CCM, we index by s the subpopulation in question. Then (15) holds if we subscript prices by s, and, for any two households m and n in the same subpopulation s,

$$\begin{aligned}\Delta \ln[\partial U_t(c_{mt}, l_{mt})/\partial c_{mt}] &= \ln(\lambda_{t+1s}) - \ln(\lambda_{ts}) \\ &= \Delta \ln[\partial U_t(c_{nt}, l_{nt})/\partial c_{nt}]. \quad (28)\end{aligned}$$

From (28) one can see that a more limited set of restrictions can be imposed in the estimation procedure. In this case S sets of time dummies must be estimated per year, where S is the number of subpopulations, instead of just one set as would be the case if the economy were fully integrated. We exploit this smaller set of identifying restrictions in our empirical analysis by allowing the state-contingent prices and the spot prices between Germany and the United States to differ.

2. APPLICATION

2.1 Parameterizing the Model

To consider a two-country model, we now index goods by their location, s, and assume that agents living in country s care only about goods that are available there. In this model investigating whether markets are integrated reduces to testing whether trade across borders is free, transport costs are negligible, and artificial barriers do not exist. In addition, the same types of physical goods are available in both countries in our model. We assume that there may be systematic differences between the two countries that are reflected by country-specific parameters in the underlying utility functions and technology. In this application, we use the following notation: l_{0nts} is female leisure, l_{1nts} is male leisure, h_{nts} is the consumption of housing services, and x_{nts} is a vector of household-specific characteristics. Let $\lambda_t = (\lambda_{t1}, \lambda_{t2})$ be the price vector for contingent claims of a unit of the consumption good. Furthermore, r_{ts} and w_{ts} are prices for housing services and labor, respectively.

We assume that the utility of the household, $U_{ts}(\cdot)$, can be expressed as the sum of two terms. The first term, $U_{0ts}(\cdot)$, reflects the preferences over housing services, and the second term, $U_{1ts}(\cdot)$, represents the preferences for male leisure. Current utility can depend on other goods and services, but we assume that it is additively separable between housing and male leisure. The choice of this specification can be justified by appealing to the existence of a home production function that depends on male leisure, female leisure, and housing services. Utility functions of this type have been used extensively in the applied econometric literature. This facilitates the comparison of the results of our study to those of previous studies. More specifically, consider the following specification of our utility function:

$$U_{ts}(l_{0nts}, l_{1nts}, h_{nts}, x_{nts}) = U_{0ts}(l_{0nts}, h_{nts}) + U_{1ts}(l_{0nts}, l_{1nts}), \quad (29)$$

where

$$\begin{aligned} U_{0ts}(l_{0nts}, h_{nts}) &= \beta^t[\alpha_{0s}^{-1} \exp(x'_{nts}B_{0s} + \varepsilon_{0nts}) h_{nts}^{\alpha_{0s}} l_{0nts}^{\alpha_{2s}}] \\ U_{1ts}(l_{0nts}, l_{1nts}) &= \beta^t[\alpha_{1s}^{-1} \exp(x'_{nts}B_{1s} + \varepsilon_{1nts}) l_{1nts}^{\alpha_{1s}} l_{0nts}^{\alpha_{3s}}], \end{aligned}$$

where B_{js} is a parameter vector to be estimated and x_{nts} is a vector of characteristics of the household (such as household size and the number of children). We do not treat female leisure as a choice variable in our framework. One of the referees correctly noted that, if we used the same framework to address female labor supply, our model would lead to stochastic singularities. This problem can be solved by introducing another term in the utility function. For example, we could add a third term of the form

$$U_{2ts}(l_{0nts}) = \beta^t[\alpha_{4s}^{-1} \exp(x'_{nts}B_{4s} + \varepsilon_{4nts}) l_{0nts}^{\alpha_{4s}}]. \quad (30)$$

Note that this term only affects the first-order condition for female labor supply and hence does not affect our estimation strategy at all because we do not exploit this condition in the estimation procedure. At the same time, it adds another stochastic error term that eliminates stochastic singularities. The wage equation is given by

$$w_{nts} = w_{ts} \exp(x'_{nts}B_{2s} + \varepsilon_{2nts}). \quad (31)$$

The exponential term captures the idea that agents with different characteristics have different productivities, and w_{ts} is the equilibrium wage for one efficiency unit of labor. The vector of characteristics influencing the labor productivity include education and work experience. For the rent equation, we assume that the rent, r_{nts}, can be expressed as the product of the number of housing services of the housing unit, h_{nts}, times the price for a standard unit of housing services, r_{ts}:

$$r_{nts} = r_{ts} h_{nts}. \quad (32)$$

Because housing services are not observed by the econometrician, we assume that housing services admit a decomposition into services that can be attributed to characteristics of the housing unit. The characteristics of the housing unit, x_{nts}, are described in our analysis by the number of rooms and an indicator for the location of the housing unit,

$$h_{nts} = \exp(x'_{nts}B_{3s} + \varepsilon_{3nts}). \quad (33)$$

Hence, we obtain the following rent equation:

$$r_{nts} = r_{ts} \exp(x'_{nts}B_{3s} + \varepsilon_{3nts}). \quad (34)$$

The vector of disturbances $(\varepsilon_{0nts}, \varepsilon_{1nts}, \varepsilon_{2nts}, \varepsilon_{3nts})$ is assumed to be identically and independently distributed across n but not necessarily t.

2.2 Orthogonality Conditions

In this particular application the orthogonality conditions come from four sources. First is the marginal rate of substitution (MRS) between housing and male labor supply, which holds period by period. [Note that this approach only exploits interior solutions of the household's optimization problem. For a treatment of corner solutions in a life-cycle model, see Heckman and MaCurdy (1980). More recently, Altug and Miller (1994) estimated a model of female labor-market participation with endogenous experience.] It does not rely on any intertemporal smoothing through financial markets. Second is the MRS between male labor supply in different periods. Substituting the parametric specification of the utility function into (15) and taking logarithms yields the following intertemporal MRS of male leisure:

$$\begin{aligned} e_{0nts}(\theta) &= \Delta x'_{nts}B_{1s} + (\alpha_{1s} - 1)\Delta \ln(l_{1nts}) \\ &\quad + \alpha_{3s}\Delta \ln(l_{0nts}) - \ln(\beta) - \Delta \ln(\lambda_t) \\ &\quad - \Delta \ln(w_{nts}), \end{aligned} \quad (35)$$

where $e_{0nts}(\theta_0) = -\Delta\varepsilon_{1nts}$. Using the same procedure to evaluate the MRS between housing and male leisure, we obtain

$$\begin{aligned} e_{1nts}(\theta) &= x'_{nts}(B_{0s} - B_{1s}) + (\alpha_{0s} - 1)x'_{nts}B_{3s} \\ &\quad - (\alpha_{1s} - 1)\ln(l_{1nts}) + (\alpha_{2s} - \alpha_{3s})\ln(l_{0nts}) \\ &\quad - \ln(r_{ts}) + \ln(w_{nts}), \end{aligned} \quad (36)$$

where $e_{1nts}(\theta_0) = \varepsilon_{1nts} - \varepsilon_{0nts} - (\alpha_{0s} - 1)\varepsilon_{3nts}$. Further orthogonality conditions can be obtained from the wage equation. Taking logarithms of Equation (31) yields

$$e_{2nts}(\theta) = \ln(w_{nts}) - x'_{nts}B_{2s} - \ln(w_{ts}), \quad (37)$$

where $e_{2nts}(\theta_0) = \varepsilon_{2nts}$. For the rent equation, we obtain

$$e_{3nts}(\theta) = \ln(r_{nts}) - x'_{nts}B_{3s} - \ln(r_{ts}), \quad (38)$$

where $e_{3nts}(\theta_0) = \varepsilon_{3nts}$.

2.3 The Estimation Procedure

As shown previously, the model leads to four testable orthogonality conditions, Equations (35)–(38), for each country. For notational convenience, we express each orthogonality condition as $e_{jnts}(\theta_0)$, where θ_0 is the unknown true parameter vector. The orthogonality conditions directly imply that

$$E[e_{jnts}(\theta_0)] = 0. \quad (39)$$

Additional orthogonality conditions can be formed by using instruments z_{jnts}, which satisfy the restriction

$$E[z_{jnts} \otimes e_{jnts}(\theta_0)] = 0. \tag{40}$$

Let

$$f_n(\cdot, \theta) = \left\{ \begin{array}{c} z_{1n11} \otimes e_{1n11}(\theta) \\ \cdots \\ \cdots \\ z_{4nT2} \otimes e_{4nT2}(\theta) \end{array} \right\}. \tag{41}$$

Following Hansen (1982), a generalized method of moments estimator can then be defined as

$$\theta_N = \arg\min_{\theta \in \Theta} \left\{ \left[\frac{1}{N} \sum_{n=1}^{N} f_n(\cdot, \theta) \right]' W_N \left[\frac{1}{N} \sum_{n=1}^{N} f_n(\cdot, \theta) \right] \right\}, \tag{42}$$

where W_N is a positive definite weighting matrix. A standard two-stage procedure was used. The first round uses the identity matrix as the weighting matrix to obtain a consistent estimate of θ. The first-round parameter estimate is then used to obtain an optimal weighting matrix, which in our model is any consistent estimator of

$$\Sigma^{-1} = E[f(\theta_0)f(\theta_0)']^{-1}. \tag{43}$$

The second stage reestimates the model with the optimal weighting matrix. Under standard assumptions, the estimator is consistent and $N^{1/2}(\theta_N - \theta_0)$ converges to a normal random vector with covariance matrix $(D'\Sigma^{-1}D)^{-1}$, where $D = E[\partial f(\cdot, \theta)/\partial \theta_0]$.

3. DATA

Our data are an extract of the PSID (waves 1980–1989) and the GSOEP (waves 1984–1991). The GSOEP is the only nationally representative panel study of households and individuals in the Federal Republic of Germany, and its content is similar to the PSID. Both datasets consider all adult members of sample households to be members of the original sample, along with their children who become adults during the course of the panel. Interviews are conducted in both studies once a year in late spring and early summer. The types of survey instruments, the structure of the questionnaires, and the variable definitions are also similar.

This study focuses on the behavior of households consisting of married couples for the purpose of comparability both between countries and with previous findings. Household records in both samples have data on the following variables: Annual hours of leisure by the husband, l_{1nts}, and the wife, l_{0nts}; the real average hourly earnings of the husband, w_{nts}; real annual rent payment for the housing unit, r_{nts}; the number of household members; the number of children (under 16); the husband's education and labor-market experience; the number of rooms in the housing unit; an indicator for the size of the city where the household lives; and a regional dummy variable for the households in the PSID. Appendix B describes how the variables used in our analysis were constructed and the requirements that were imposed. There are 350 German couples about whom information was collected from 1983 to 1990 and 499 couples from the United States, where the sample phase is 1980–1988.

Descriptive statistics of the sample extracts are given in Tables 1 and 2. On average, household heads in our PSID subsample are roughly eight years older than those in the German sample. The average age of a German household head was 40 years in 1984, whereas the average age for an American household head was 48 in 1980. The average household size in the PSID sample is 3.9, significantly bigger than the corresponding number for the German sample, 3.5. This fact is also reflected in the average number of children per household.

On average, an American household lives in an apartment that has significantly more rooms than the apartment of an average German household. There are at least three possible explanations. Americans may prefer bigger apart-

Table 1. Descriptive Statistics for the GSOEP Subsample

	Year							
Variables	*1983*	*1984*	*1985*	*1986*	*1987*	*1988*	*1989*	*1990*
Household size	3.45	3.48	3.50	3.53	3.53	3.53	3.53	3.49
	(1.06)	(1.04)	(1.06)	(1.07)	(1.04)	(1.06)	(1.09)	(1.08)
Number of children under 16	1.10	1.09	1.04	1.03	1.01	.96	.93	.89
	(.97)	(.99)	(1.02)	(1.05)	(1.05)	(1.08)	(1.08)	(1.07)
Number of rooms	4.14	4.19	4.19	4.14	4.16	4.19	4.16	4.18
	(1.38)	(1.36)	(1.40)	(1.38)	(1.40)	(1.43)	(1.39)	(1.41)
Rent[a]	—	705.92	745.59	787.36	813.38	850.08	924.48	1,002.57
	—	(346.91)	(352.71)	(388.56)	(409.85)	(413.93)	(456.43)	(486.23)
Hours worked$^{b}_{m}$	44.28	44.38	44.26	43.62	43.73	43.83	43.19	43.29
	(8.31)	(6.62)	(7.36)	(7.40)	(7.45)	(6.61)	(5.91)	(6.53)
Gross labor income$^{a}_{m}$	3,578.63	3,693.23	3,936.23	4,122.32	4,247.84	4,434.63	4,623.92	4,884.37
	(1,170.48)	(1,254.28)	(1,526.49)	(1,632.20)	(1,622.73)	(1,621.31)	(1,712.98)	(1,927.18)
Hours worked$^{b}_{f}$	30.40	27.45	28.21	28.44	27.91	27.73	26.99	25.90
	(14.29)	(12.81)	(13.41)	(11.61)	(13.21)	(13.11)	(13.29)	(12.89)

NOTE: Standard errors are given in parentheses.
[a] Measured on a monthly basis in DM.
[b] Measured on a weekly basis in DM.
$_m$ Variable refers to male.
$_f$ Variable refers to female.

Table 2. Descriptive Statistics for the PSID Subsample

	Year								
	1980	1981	1982	1983	1984	1985	1986	1987	1988
Household size	3.98	3.92	3.87	3.82	3.84	3.83	3.78	3.83	3.82
	(1.65)	(1.60)	(1.52)	(1.40)	(1.42)	(1.42)	(1.26)	(1.25)	(1.23)
Number of children	1.57	1.55	1.54	1.53	1.58	1.63	1.64	1.65	1.65
under 16	(1.34)	(1.33)	(1.28)	(1.24)	(1.25)	(1.24)	(1.20)	(1.19)	(1.17)
Number of rooms	5.40	5.42	5.56	5.52	5.59	5.61	5.59	5.69	5.91
	(1.53)	(1.54)	(1.49)	(1.45)	(1.53)	(1.51)	(1.53)	(1.64)	(1.71)
Rent[a]	221.03	241.85	266.84	273.22	301.70	322.37	334.87	—	—
	(110.76)	(126.11)	(139.86)	(127.97)	(147.72)	(161.70)	(166.73)	—	—
Hours worked$^{b}_{m}$	42.76	41.48	40.57	41.97	42.85	42.86	43.33	43.96	43.72
	(12.10)	(12.15)	(12.00)	(12.13)	(11.70)	(11.37)	(11.86)	(11.58)	(11.88)
Gross labor income$^{a}_{m}$	1,425.05	1,550.87	1,605.56	1,733.35	1,886.44	1,957.76	2,052.69	2,234.64	2,334.64
	(838.74)	(946.49)	(1,007.32)	(1,072.58)	(1,337.93)	(1,198.64)	(1,247.06)	(1,456.30)	(1,391.11)
Hours worked$^{b}_{f}$	26.52	25.66	26.07	26.17	27.73	27.59	28.34	28.87	29.07
	(14.67)	(13.78)	(14.51)	(14.46)	(14.76)	(14.23)	(12.99)	(13.37)	(13.14)

NOTE: Standard errors are given in parentheses.
[a] Measured on a monthly basis in U.S. dollars.
[b] Measured on a weekly basis in U.S. dollars.
$_m$ Variable refers to male.
$_f$ Variable refers to female.

ments and houses than Germans because American have larger households. We expect that an additional member of the household increases the valuation of housing services both of American and German households. Alternatively, these differences in housing consumptions may be caused by fundamental differences in the valuation of housing services. These differences should be reflected in the estimates of α_0, the coefficient of the utility function that measures the valuation of housing services. Finally, the price of housing might be cheaper in the United States. We will return to these issues in Section 4. Rental expenditures for houses and apartments rose throughout the observation period in Germany, especially at the end of the observation period. Nominal rents also increased throughout the observation period in the United States, although these increases are in line with the general inflation that occurred throughout the decade.

Whereas labor supply, as measured by the average number of hours per week, showed an increasing trend throughout the 1980s for the American sample, the opposite is true for the German sample. During the period from 1983 to 1988, for which we have data on both countries, male labor supply increased in the PSID sample by 4.2% but decreased by 1.0% in the German sample. Some of these changes in the labor supply can be explained by the cyclical behavior of the two economies. The American economy went through a cyclical downturn in the beginning of the 1980s. The increase in labor supply during the period from 1983 to 1988 parallels the increase in economic activity in the United States. The recession in the early 1980s was much less severe in Germany. Whereas the American economy showed a strong cyclical pattern throughout the 1980s, the German economy grew at modest rates between 1.0% and 2.5% throughout this time period. These rates were considered to be too low to sustain a substantial increase in the work force. Our model produces estimates of the realized aggregate shocks in the economy. These estimated shocks should reflect the general cyclical trends of the two economies. Section 4 shows how our model can be used to explain the observed behavior in both countries.

The general trends of both economies are also reflected in the wage data. Real wages stagnated in the United States throughout the 1980s. In the PSID sample, the real wage of males in 1979 and in 1985 was almost identical, $10.48 per hour. Real wages grew moderately in Germany. The average growth rate over the eight-year period of the sample was 1.58%. The framework includes estimates of wage functions for both countries. This allows us to decompose changes in the real wage into changes due to general levels of education and experience and changes that are induced by increases in the compensation for an efficiency unit of labor. Alternatively, we might expect that the relative productivity of one economy increased relative to the other economy. This increase in productivity should be reflected in the real exchange rate between the United States and Germany. The framework presented in this article allows us to estimate commodity-specific exchange rates, at least for the labor market. Section 4 tests whether the commodity-specific exchange rate for an efficiency unit of labor changed throughout 1983 to 1990, the period for which we have observations from both datasets.

4. EMPIRICAL RESULTS

Except where indicated, the same instruments were used throughout our estimation. The components of our instrument vector included a constant, twice-lagged values of the logarithm of male and female leisure, the household size, the number of children, education, education squared, experience, and experience squared. Our empirical findings are summarized in Tables 3, 4, and 5. (First, we estimated the model separately for the two datasets. Because the results hardly differ from those we obtained in joint estimation, they are not reported here.) Note that the reported standard errors in the tables refer to the asymptotic covariance matrix of $N^{1/2}(\theta_N - \theta_0)$. To calculate standard t tests, these

Table 3. Estimation of the MRS Functions

Parameters of utility function	Variable	I SOEP	I PSID	II SOEP	II PSID	III SOEP	III PSID	IV SOEP	IV PSID
$\alpha_0 - 1$		−2.02	−2.08	−4.58	−2.32	−4.19	−1.91	−3.17	−.91
		(.22)	(1.13)	(.24)	(.59)	(.26)	(.38)	(1.46)	(1.00)
$\alpha_1 - 1$		−1.87	−2.15	−2.46	−2.53	−3.88	−1.95	−3.76	−1.83
		(.93)	(2.81)	(1.07)	(1.68)	(.73)	(1.22)	(1.51)	(1.97)
$\alpha_2 - \alpha_3$		−.92	−1.23	−.83	−1.74	—	—	—	—
		(.89)	(2.81)	(1.04)	(1.67)				
α_2		—	—	—	—	−.31	−.66	−.29	−.61
						(.84)	(2.14)	(3.00)	(4.20)
α_3		—	—	—	—	2.09	.47	2.29	.56
						(.39)	(2.11)	(2.19)	(2.58)
ΔB	Household size	−.21	−.16	−.46	.50	—	—	—	—
		(.20)	(.26)	(.21)	(.36)				
	Number of children	.07	.10	.29	−.15	—	—	—	—
		(.22)	(.30)	(.12)	(.29)				
B_0	Household size	—	—	—	—	.33	.39	.40	.15
						(.26)	(.48)	(.58)	(1.01)
	Number of children	—	—	—	—	−.09	−.11	−.23	−.05
						(.20)	(.53)	(.81)	(.96)
B_1	Household size	—	—	—	—	.02	.13	.13	.11
						(.21)	(.31)	(.51)	(.44)
	Number of children	—	—	—	—	.22	−.15	−.13	−.12
						(.17)	(.46)	(.64)	(.42)
B_3	Size of housing unit	.28	.19	.47	.27	.48	.37	.41	.27
		(.03)	(.08)	(.02)	(.06)	(.02)	(.06)	(.11)	(.16)
	City indicator	.76	.52	−.11	.63	.94	.55	1.37	.54
		(.05)	(.24)	(.04)	(.15)	(.05)	(.02)	(.39)	(.23)
	J value	197.74		333.48		331.46		242.42	
	Degrees of freedom	207		319		321		209	

NOTE: Standard errors are given in parentheses. To calculate t tests, divide coefficients by estimated standard errors and multiply result by $N^{1/2}$. Only five instruments were used in IV.

standard errors should be divided by $N^{1/2}$, which is approximately 20 in our application. Table 3 presents results obtained by estimating various combinations of the four orthogonality conditions. In the first column, labeled I, we only use the MRS between housing and male leisure and the rent equation. In this setup, only a subset of the un-

Table 4. Testing Restrictions Using the MRS Functions

Parameters of utility function	Variable	I SOEP	I PSID	II SOEP	II PSID	III SOEP	III PSID	IV SOEP/PSID
$\alpha_0 - 1$		−4.19	−1.91	−5.62	−4.05	−3.60	−3.77	−2.59
		(.26)	(.38)	(.40)	(.82)	(.19)	(1.55)	(.08)
$\alpha_1 - 1$		−3.89	−1.96	−3.30	−3.49	−2.46	−4.46	−4.66
		(.73)	(1.22)	(.85)	(1.87)	(.66)	(2.10)	(.59)
α_2		−.31	−.66	.81	.34	.33	.31	−1.67
		(.84)	(2.14)	(.78)	(2.45)	(.74)	(4.45)	(.75)
α_3		2.10	.47	2.68	2.74	1.04	3.93	2.24
		(.39)	(2.11)	(.39)	(1.21)	(.83)	(3.77)	(.50)
B_0	Household size	.33	.39	.10	.28	−.63	1.06	−.11
		(.26)	(.48)	(.27)	(.62)	(.26)	(.74)	(.27)
	Number of children	−.09	−.11	.42	−.03	.16	−.78	.77
		(.20)	(.53)	(.18)	(.74)	(.18)	(.86)	(.20)
B_1	Household size	.03	.13	.29	−.37	.59	.36	−.93
		(.21)	(.31)	(.20)	(.34)	(.18)	(.61)	(.24)
	Number of children	.22	−.15	.18	.27	−.46	−.58	.97
		(.17)	(.46)	(.16)	(.47)	(.15)	(.75)	(.19)
B_3	Size of housing unit	.48	.37	.26	.25	.39	.19	.26
		(.03)	(.06)	(.01)	(.06)	(.02)	(.05)	(.01)
	City indicator	.94	.55	.49	.71	1.27	.13	.90
		(.05)	(.02)	(.04)	(.01)	(.07)	(.01)	(.03)
	J value	331.34		341.24		318.56		354.79
	Degrees of freedom	321		325		313		335

NOTE: Standard errors are given in parentheses. To calculate t tests, divide coefficients by estimated standard error and multiply result by $N^{1/2}$.

Table 5. Estimation of the Wage Functions

		I		II		III	
Parameters	*Variable*	*SOEP*	*PSID*	*SOEP*	*PSID*	*SOEP*	*PSID*
	Education	.132	.132	.095	.086	.134	.134
		(1.13)	(.15)	(1.06)	(.15)	(.122)	(.122)
	Education squared	−.0013	−.0012	−.0001	.0009	−.0021	−.0021
		(.040)	(.043)	(.037)	(.0044)	(.0043)	(.0043)
	Experience	.020	.016	.014	.014	.0094	.0094
		(.096)	(.041)	(.096)	(.041)	(.034)	(.034)
	Experience squared	−.0002	−.0002	−.0001	−.0001	−.00008	−.00008
		(.0017)	(.0006)	(.0005)	(.0017)	(.0005)	(.0005)
$\ln(w_{80})$		—	.62	—	.87	—	.81
			(1.70)		(1.71)		(1.15)
$\ln(w_{81})$		—	.61	—	.86	—	.81
			(1.69)		(1.70)		(1.16)
$\ln(w_{82})$		—	.58	—	.83	—	.77
			(1.72)		(1.73)		(1.17)
$\ln(w_{83})$		—	.59	—	.84	—	.78
			(1.90)		(1.92)		(1.59)
$\ln(w_{84})$		—	.55	—	.80	—	.76
			(1.78)		(1.80)		(1.19)
$\ln(w_{85})$		1.14	.53	1.50	—	1.47	.74
		(7.95)	(1.79)	(7.40)		(1.28)	(1.20)
$\ln(w_{86})$		1.14	.57	1.51	—	1.47	.78
		(7.95)	(1.78)	(7.40)		(1.26)	(1.20)
$\ln(w_{87})$		1.15	.60	1.53	—	1.48	.82
		(7.96)	(1.77)	(7.41)		(1.50)	(1.20)
$\ln(w_{88})$		1.17	.63	1.55	—	1.50	.86
		(7.97)	(1.77)	(7.41)		(1.27)	(1.21)
$\ln(w_{89})$		1.19	—	1.57	—	1.53	—
		(7.98)		(7.43)		(1.27)	
$\ln(w_{90})$		1.19	—	1.59	—	1.54	—
		(7.99)		(7.43)		(1.28)	
$\ln(\varepsilon_{85-89})$		—	—	—	−.70	—	—
					(7.71)		
	J value	127.19		130.29		137.47	
	Degrees of freedom	112		115		116	

NOTE: Standard errors are given in parentheses. To calculate *t* tests, divide coefficients by estimated standard error and multiply result by $N^{1/2}$.

known parameters is identified, more specifically the difference of the parameters that affect the marginal utility of male leisure and housing. Note that our estimates for the utility parameters of housing and male leisure, α_0 and α_1, have the expected sign. They are both significantly smaller than 1, indicating that the objective function is concave in these parameters as required by the theory.

The parameters in the rent function, B_3, also have the expected sign. The coefficient for the size of the housing unit is positive. Clearly, bigger apartments and houses provide a higher utility than smaller ones. The coefficient for Germany is bigger than the one for the United States. The city indicator is positive, suggesting that apartments in the city are preferred to those in the country. A drawback of the estimation of the rent function is that the estimates for the prices of a standard unit of housing are very high. The variables that we use to construct the housing services variable, size and city indicator, do not explain a large portion of the variation in the data. This problem could be partially solved by including more variables in the rent regression. The GSOEP provides a variety of additional information on the quality of the housing unit, but the PSID does not. Because we limited the analysis to variables that are common to both datasets, only the two mentioned in the text were included.

Comparing the estimates for the coefficients across the two countries, we find that there is a surprisingly high degree of similarity between them. Most of the coefficients have not only the same sign but appear to be reasonably close in magnitude. The J statistic for the overidentifying restrictions is given by

$$J_N = N\left\{\left[\frac{1}{N}\sum_{n=1}^{N} f_n(\cdot,\theta_N)\right]' W_N \left[\frac{1}{N}\sum_{n=1}^{N} f_n(\cdot,\theta_N)\right]\right\}. \tag{44}$$

As shown by Sargan (1958) and Hansen (1982) the test statistic, J_N, converges in distribution under the null hypothesis to a χ^2 distribution with j df, where j is the number of overidentifying restrictions (number of orthogonality conditions − number of parameters). The test for the overidentifying restrictions is not rejected at standard levels of significance.

The second column, labeled II, adds the wage equation to the set of orthogonality conditions. Comparing the parameter estimates of column II with column I, we find that adding additional orthogonality conditions decreases the estimates for α_0 and α_1 for both the American and the German sample. In addition, the coefficients for the size of the

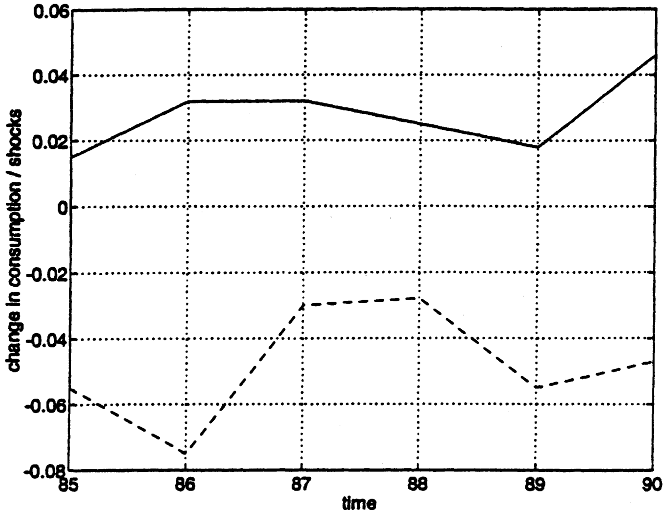

Figure 1. Aggregate Consumption and Shocks in Germany: ———, Change of Aggregate Real Consumer Expenditures; – –, Estimated Change of Price Realizations for Contingent Claims.

housing unit increase in both samples. The coefficient for the city indicator is negative for the German sample, which is slightly disturbing, although this is the only counterintuitive sign we found. The standard errors obtained in this estimation are lower than their counterparts in column I, illustrating the point that adding orthogonality conditions typically increases the precision of an estimator.

To identify all the structural parameters of interest, the orthogonality conditions that characterize the intertemporal marginal rate of substitution must be included in the estimation framework. Column III presents the results of the estimation when we use orthogonality conditions from Equations (35)–(37). The estimates for α_0 and α_1 are roughly the same as in the first two columns. Turning to female leisure, the estimates for α_2 and α_3 range from $-.66$ to 2.09, less than 1 in three out of four cases. Because female leisure is not a choice variable, there are no grounds for requiring household utility to be concave in that argument. In an extended analysis in which female labor supply was

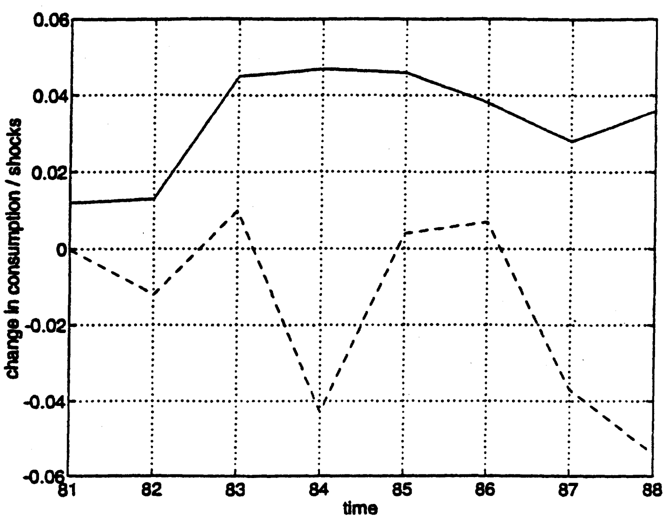

Figure 2. Aggregate Consumption and Shocks in the United States: ———, Change of Aggregate Real Consumer Expenditures; – –, Estimated Change of Price Realizations for Contingent Claims.

endogenous, the findings from this study show it should enter the utility nonseparably with housing, male leisure, and other variables as well.

Parameters that affect the marginal utilities of housing and leisure, B_0 and B_1, are also identified from the orthogonality conditions supporting column III. The coefficients for the household size are positive, meaning that an additional family member increases the marginal utility of housing and male leisure. The coefficients for the number of children are negative in three out of four cases, which indicates that children have a smaller impact than adults on housing and labor supply.

Column IV of Table 3 presents the parameter estimates from deploying all four orthogonality conditions jointly in the estimation. Here a smaller instrument vector was used, and this may explain the increase in estimated standard errors. The original set of instruments proved unwieldy, with more than 500 orthogonality conditions. The estimation program encountered inversion problems when calculating the optimal weighting matrix. Hence, we only used a constant and twice-lagged values of male and female leisure, household size, and the number of children as instruments in this part of the estimation. Comparing column IV with columns I–III, we find that the estimates for the α's of the utility function increase slightly. The coefficients for the household size are positive in all four cases, whereas the coefficients for the number of children are negative. Children have a small and insignificant effect on male labor supply. (Note that the impact of an additional child is measured by the sum of the coefficients for household size and number of children.) On the other hand, an extra child has a positive effect on the demand for housing, with a magnitude almost as big as an adult.

Under CCM, aggregate shocks are fully transmitted to household choices through the prices of contingent claims. This implies that the estimates of the price realizations of the contingent claims can be interpreted as measures for aggregate shocks. The estimation procedure treats the logarithm of the price ratios as dummy variables. To see how well aggregate fluctuations are explained by our model, we plotted in Figures 1 and 2 the changes of real consumer expenditure and the estimated changes of price realizations of the contingent claims against time. Positive aggregate shocks should be reflected by decreasing prices for contingent claims.

The results are remarkable. Consider Figure 1, Germany. The estimated changes of the price realizations are all negative, which implies that prices for contingent claims decreased throughout the seven years of the observation period in Germany. These price decreases were paralleled by steady increases in aggregate household expenditures. Our measure of aggregate shocks captures the aggregate behavior of the German economy rather well. Between 1984 and 1986 the German economy grew at moderate rates, then slipped into a stagnation period between 1986 and 1988, and experienced a new expansion in 1989 and 1990 due to the unification process.

Figure 2 shows the same graph for the United States. Again, our estimates pick up the cyclical behavior of the

U.S. economy in the 1980s reasonably well. The price estimates for the contingent claims reflect both the recession in the early 1980s and the boom periods later in the decade. Note that aggregate consumption grew much faster than prices declined, indicating that some growth in consumer expenditure may have occurred because of an inflow of foreign capital during the Reagan administration rather than improved economic efficiency.

Summarizing these results, our model seems to perform well along several dimensions. None of its versions are rejected by the overidentifying restrictions, and the associated test statistics are well below their critical values. The parameter values for housing and male leisure, α_0 and α_1, have the expected signs and are of a reasonable magnitude. The estimated coefficients for the rent function are slightly more problematic. Nevertheless, the estimates for the size and the location indicator have the expected signs in almost all cases. Most of the other coefficients are highly significant and have signs and magnitudes comparable with previous studies (e.g., see Altug and Miller 1990; Ham and Jacobs 1994). In addition, the estimates of the parameters of the utility functions do not differ drastically across the two datasets. The estimates for the realizations of the contingent price processes capture aggregate shocks over the sample period.

Because our model performs reasonably well along the dimensions outlined previously, we then used it as a benchmark to test hypotheses about the degree of market integration between the two countries. The results of these tests, which impose various restrictions on our framework, are presented in Tables 4 and 5. Let J_N^U be the J statistic for the unrestricted model and J_N^R be the one for the restricted model. The validity of these restrictions can be tested using a D statistic, which is defined by $D_N = J_N^R - J_N^U$. Gallant and Jorgenson (1979) and Newey and West (1987) showed that D_N converges in distribution under the null hypothesis to a χ^2 distribution with d df, where d is the number of restrictions imposed on the parameter space.

First, we address the question of whether the prices of contingent claims of the numeraire good are equal between the countries. Because we have four overlapping time periods between the two samples, this test imposes four restrictions on the parameter vector of our model. The results of this test are presented in the second column of Table 4. The value of the test statistic is approximately 9.9, implying that the data rejects this hypothesis at the .05 level. This finding shows that markets are not completely integrated across the two countries. That is, despite appearances to the contrary, national boundaries still play an important role in the determination of resource allocations. As mentioned earlier, the prices for the contingent claims can also be interpreted as measures for aggregate shocks in the economy. From this perspective, the results indicate that aggregate shocks are not identical for the two countries. Country-specific components still play a major role in national economies. It should, however, be pointed out that the general test for all overidentifying restrictions is not rejected at standard levels of significance, and imposing the assumption of a single CCM economy does not contort the data.

At the suggestion of a referee, we also tested for comparison purposes whether markets are fully integrated within the United States. This test was implemented by partitioning the United States in two separate ways (North–South and East–West). Supposing that markets are complete in both geographical regions but that the prices for contingent claims might differ between the two regions, there are three sets of prices for the contingent claims, $\lambda_t = (\lambda_{t0}, \lambda_{t11}, \lambda_{t12})$. For every household in the PSID, define a dummy variable, d_{nt}, which is equal to 1 if household n lives in the South in periods t and $t+1$ and 0 otherwise. Using the terminology of Section 2.4, we can express Equation (28) for a household that did not move between the two regions in period t and $t+1$ as

$$\Delta \ln[\partial U(c_{nt}, l_{nt})/\partial l_{nt}] \\ = d_{nt}\Delta \ln(\lambda_{t11}) + (1 - d_{nt})\Delta \ln(\lambda_{t12}). \quad (45)$$

This generalization of the model adds eight additional parameters to be estimated. The results of the estimation of the model that divides the PSID into north and south observations are reported in the third column of Table 4. (We also tested the case in which the United States is divided into an eastern and a western region. But because the results are similar to the ones reported previously, we do not discuss them here.) A comparison between the first and the third columns of Table 4 shows that the estimated parameter values are similar in both cases. The test statistic for the hypothesis that the two sets of contingent claim prices for the United States are equal is 12.78, which is below the critical value of 13.36 for a .1-level test. This finding suggests that there is more integration between different regions in the United States than between the United States and Germany.

A surprising finding of our analysis is that the utility parameter estimates for the German and the American subsamples are quite similar, surprising because the institutional differences between the two economies are quite marked. For example, wage determination is more decentralized in the United States than in Germany. If these institutional or cultural differences had been important, then we might have expected to pick up bigger differences in the estimated coefficients of our model. Continuing this line of investigation further, we tested the other extreme to see whether the hypothesis that the parameters of the utility functions are equal across the two countries can be rejected. The results of this test are presented in the fourth column of Table 4. The value of the test statistic is 23.45, the restriction is rejected at the 90% confidence level but fails to reject at the 95% confidence level. Thus, preferences exhibit a significant country-specific component, and allowing for such differences seems to improve the estimation results. Their overall importance is not critical, however; for example, the overidentifying restrictions of the restricted model (i.e., taken as a separate entity rather than being compared with the model that allows country-specific utility functions) are not rejected, and the parameter estimates obtained remain reasonable.

The last part of our report considers the results obtained from estimating wage functions. Because the parameters of the wage functions do not appear in the other three orthogonality conditions, we report the results for the estimation of the wage function separately in Table 5. The first column shows the results for an unrestricted estimation of the wage function. The signs of the estimates are as expected, and the coefficients for the two countries are close to each other. The productivity of the agents is a concave function in experience and education.

The estimates for the returns of an efficiency unit of labor are significant over time for both samples. The real wage, measured by the price of an efficiency unit, increased moderately during the sample period in Germany, approximately 5% over the six-year period. The real wage in the United States followed a more cyclical pattern, decreasing in the beginning of the 1980s and increasing over the second part of the decade which reflects the widespread opinion that real wages had been stagnating in the United States throughout the 1980s (see Davis 1992; Katz and Murphy 1992).

An appealing feature of our model is that price equalization across countries can be tested. Define the real exchange rate for an efficiency unit of labor to be

$$e_t^{2,1} = \frac{w_{1t1}}{w_{1t2}}. \tag{46}$$

The Purchasing Power Parity Theorem (PPP) states that the real exchange rate is constant over time. This is the most common version of PPP. A more recent approach interprets PPP as a long-run neutrality proposition: Long-term effects of monetary policy leave the real exchange rate unaffected. Clearly, our framework is silent on monetary policy. [For a recent discussion of PPP, see McCallum (1995).] The results of testing this proposition are reported in the second column of Table 5. We obtained an estimate of 2.01 for the exchange rate over the four-year period from 1985 to 1988. The value of the test statistic is approximately 3.1, which compares with a critical value of 7.81 for a .05-level test. Although our test is limited to a short time span, we find no evidence against PPP.

This test of PPP differs from previous tests, which have used time series data on prices, exchange rates, and interest rates. [Some recent studies were done by MacDonald (1993), Johansen and Juselius (1992), and Abuaf and Jorion (1990).] Our approach estimates prices from a wage regression and treats the price of an efficiency unit of labor as an unknown parameter to be estimated. Much of the attention in the recent literature on PPP focuses on whether the time series are stationary, whereas the asymptotic distribution of our estimator is obtained by averaging over the number of households in the sample, N, not by averaging over the number of time periods, T. In this way our approach avoids most of the pitfalls of modern time series analysis and is able to entertain nonstationary aggregates. Finally, we consider a commodity-specific exchange rate, the real exchange rate for one efficiency unit of labor, rather than the exchange rate for a bundle of consumption goods. Given these differences, it is difficult to make specific comparisons between our results and those of previous studies. Generally speaking there is a controversy within the time series literature about the relevant empirical regularities, but most of those studies show that PPP is only of limited usefulness in explaining and predicting exchange rates in the short run. Our results do not conflict with this overall assessment but indicate that relative wages do not vary much between countries over a medium horizon.

Finally, we test whether the productivities of the agents follow the same pattern in the two countries. This imposes four restrictions on the parameter vector. The last column of Table 5 presents the results of this test. The value of this test, 10.3, falls within the critical region of a .05-level test. Apparently, returns for education and experience differ significantly across the two countries, although this finding should be tempered with the qualification that age characteristics differ across the two samples. Moreover, structural differences in the educational systems of both countries and differences in the labor-market institutions might be contributing to this finding.

5. CONCLUDING REMARKS

Our empirical findings show that making international comparisons based on structural models of competitive equilibrium using panel data of households living in different countries is informative. The tests of the overidentifying restrictions do not provide any evidence against the validity of the model. In particular, the assumption of complete and competitive markets within countries is not rejected. The estimated values of the coefficients are, for the most part, similar to those obtained by others and are highly significant. The main evidence against the specifications we investigate come from two sources: A few of the coefficients take on implausible values, and the estimated values are somewhat sensitive to precisely which sets of orthogonality conditions are imposed.

Notwithstanding this caveat, some important empirical regularities emerge from our empirical investigations. First, the estimated preferences for housing consumption and male labor supply, and also the estimated wage functions, are similar across both countries. Moreover some of the differences that were found are robust to the choice of orthogonality conditions. These differences can be attributed to institutional differences between Germany and the United States. Second, despite the lack of evidence against the equality of marginal rates of substitution functions within countries, we did uncover indirect evidence against the economic integration across national boundaries in financial markets. On the other hand, real wages in both countries appear to move in lock step, thereby supporting real purchasing power parity in the labor market. Third, our estimates for the realizations of the prices of contingent claims faithfully reflect fluctuations in aggregate time series.

We view the empirical findings from this study as being quite promising for further work that combines the economic and econometric theory with panel data on households from different countries to address questions in international economics. Our results show that differences in the estimated structural parameters between countries are of a similar magnitude to the differences that emerge in the estimation using different sets of orthogonality conditions. Although there are important cultural and institutional differences between Germany and the United States, our empirical results are obtained from a framework that entirely ignores all rigidities to international trade. This suggests that these institutional barriers do not constitute an insurmountable barrier to making comparisons about the allocation of real resources in these economies at the microeconomic level.

The comparability we find between Germany and the United States contrasts with the differences other researchers have obtained when they adopt econometric practices that we criticized in Section 1. We believe that if the techniques we argue as inappropriate had been adopted here, then the apparently contradictory findings found elsewhere would be mirrored in our data as well. This also suggests that progress in the field will be hampered until the substantial disagreement about these methodological matters dissipates.

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APPENDIX A: THE CORRELATION MATRIX OF THE RESIDUALS

Table A.1 illustrates the estimated correlations of the residuals for the GSOEP based on the estimation reported in column III of Table 3.

APPENDIX B: CONSTRUCTION OF THE DATASETS

We selected households composed of couples who had been married throughout the survey period from 1984 through 1991. Furthermore, we excluded all households that did not have complete data throughout the survey period. We started out with 9,236 individuals in the GSOEP. After eliminating all households that did not consist of a married couple, we had 1,455 observations left. The losses associated with the selection criteria were as follows:

1. Missing wage data for the waves 1984–1991: 920
2. Missing rent data for the waves 1986–1991: 150
3. Missing data on size of housing unit: 3
4. Annual hours worked less than 440: 32

This left us with a sample of 350 households. For the GSOEP, we used the following variable constructions:

1. Annual male and female leisure: We took the variable average hours worked per week and multiplied it by 52 to compute the annual hours worked. The annual leisure is equal to 8,760 minus annual hours worked.
2. Real average wage: The nominal wage was calculated according to the following formula: (12*gross monthly labor income)/(annual hours). To convert nominal to real values, we used the gross national product (GNP) deflator from the Organization for Economic Cooperation and Development (OECD) dataset.
3. Real annual rent: The nominal annual rent was obtained by multiplying the monthly rent by 12. We used the actual rent for renters and imputed rent for house and apartment owners. To convert nominal to real values, we used a price deflator for the housing market that is provided in the OECD dataset.
4. Annual family income is given by average monthly income times 12.

Table A.1. Correlations of Residuals for the GSOEP

	*R*186	*R*187	*R*188	*R*189	*R*190	*R*191	*R*286	*R*287	*R*288	*R*289	*R*290	*R*291	*R*386	*R*387	*R*388	*R*389	*R*390	*R*391
*R*186	1.00	.91	.88	.85	.86	.84	.79	.75	.74	.71	.73	.74	-.04	.07	.07	-.01	.08	-.05
*R*187	.91	1.00	.94	.91	.92	.89	.72	.77	.76	.73	.76	.76	-.02	.07	.07	-.03	.09	-.05
*R*188	.88	.94	1.00	.89	.90	.88	.71	.75	.77	.71	.74	.75	-.03	.10	.07	-.02	.09	-.04
*R*189	.85	.91	.89	1.00	.92	.89	.72	.77	.75	.79	.78	.77	-.02	.07	.06	-.00	.09	-.03
*R*190	.86	.92	.90	.92	1.00	.94	.73	.77	.75	.72	.79	.78	-.08	.11	.04	-.00	.06	-.03
*R*191	.84	.89	.88	.89	.94	1.00	.71	.75	.73	.70	.74	.77	-.02	.12	.02	-.00	.07	-.02
*R*286	.79	.72	.71	.72	.73	.71	1.00	.95	.92	.90	.92	.92	-.08	.05	.12	-.04	.12	.01
*R*287	.75	.77	.75	.77	.77	.75	.95	1.00	.96	.94	.97	.96	-.01	-.07	.17	-.02	.12	.04
*R*288	.74	.76	.77	.75	.75	.73	.92	.96	1.00	.92	.95	.94	.00	-.00	.04	.02	.13	.05
*R*289	.71	.73	.71	.79	.72	.70	.90	.94	.92	1.00	.94	.93	.02	-.03	.09	-.08	.21	.03
*R*290	.73	.76	.74	.78	.79	.74	.92	.97	.95	.94	1.00	.96	-.04	-.00	.08	-.03	.08	.07
*R*291	.74	.76	.75	.77	.78	.77	.92	.96	.94	.93	.96	1.00	-.00	.01	.08	-.04	.13	-.05
*R*386	-.04	-.02	-.03	-.02	-.08	-.02	-.08	-.01	.00	.02	-.02	-.00	1.00	-.40	-.07	-.02	-.00	.12
*R*387	.07	.07	.10	.07	.11	.12	.05	-.07	-.00	-.03	-.00	.01	-.40	1.00	-.40	-.02	.01	-.15
*R*388	.07	.07	.07	.06	.04	.02	.12	.17	.04	.09	.08	.08	-.07	-.40	1.00	-.47	.00	-.04
*R*389	-.01	-.03	-.02	-.00	-.00	-.00	-.04	-.02	.02	-.08	-.03	-.04	-.02	-.02	-.47	1.00	-.44	.05
*R*390	.08	.09	.09	.09	.06	.07	.12	.12	.13	.21	.08	.13	-.00	.01	.00	-.44	1.00	-.27
*R*391	-.05	-.05	-.04	-.03	-.03	-.02	.01	.04	.05	.03	.07	-.05	.12	-.15	-.04	.05	-.27	1.00

NOTE: $R186$–$R191$: Residuals of rent equation; that is, $R186 = v_{2,n,86,1}$. $R286$–$R291$: Residuals of MRS equation; that is, $R286 = u_{1,n,86,1} - u_{0,n,86,1} - (\alpha_{0,1} - 1)v_{2,n,86,1}$. $R386$–$R391$: Residuals of *i*MRS equation, $R386 = -\Delta u_{1,n,86,1}$.

5. Labor-market experience is given by age minus 6 minus number of years in school.

6. Total education is the number of years in school plus the number of years in university, apprenticeships, vocational school, and so forth.

7. The city indicator is 1 if the city has more than 100,000 inhabitants.

8. Number of rooms, household size, and number of children under 16 are variables in the dataset.

The original PSID dataset consisted of 6,925 households. To construct our sample we eliminated all households that did not have complete data for the relevant variables, except annual rent. For the rent variable, we only required that we have at least one observation during the period from 1980 through 1987. After eliminating households that did not consist of a married couple, our sample size was reduced to 2,513. The losses associated with the remaining criteria were as follows:

1. Missing wage data 1980–1989: 675 households
2. Missing education data 1980–1989: 21 households
3. Missing experience data 1980–1989: 47 households
4. Missing rent data 1980–1987: 1,206 households
5. Missing data for size of housing unit 1980–1987: 4 households
6. Missing data for size of city 1980–1987: 14 households
7. Annual hours less than 440 1980–1989: 40 households

This left us with a sample of 499 households. The number of households for which we have rent data varies between 260 and 290. For the PSID we used the following variable constructions:

1. Annual male and female leisure: The annual leisure is equal to 8,760 minus annual hours worked.

2. Real average wage: The nominal wage was calculated according to the following formula: (annual gross labor income)/(annual hours worked). To convert nominal to real values, we used the GNP deflator from the OECD dataset.

3. Real annual rent: The PSID only provides rental information for agents who do not own their apartments and actually pay rent to a landlord. There is not an imputed rent variable for owners. To convert nominal to real values, we used a price deflator for the housing market that is provided in the OECD dataset.

4. Labor-market experience is defined as total experience in the labor market. Total education is the number of years in school plus the number of years in university, apprenticeships, vocational school, and so forth.

5. The city indicator is 1 if the city has more than 100,000 inhabitants. The two regional dummy variables were constructed by dividing the United States into two disjoint regions (East–West and North–South). The PSID includes a variable that reports the state where the household lived in a given period. The most eastern states in the West are Wisconsin, Illinois, Missouri, Arkansas, and Alabama. The most northern states in the South are Virginia, West Virginia, Kentucky, Missouri, Kansas, Colorado, Utah, Nevada, and California.

6. Annual family income, number of rooms, household size, and number of children under 16 are variables in the dataset.

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