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To See or Not To See: Another Look at Research on Temporal Trends and Cross-National Differences in Educational Homogamy

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Abstract

This paper examines a recent critical exchange over cross-national and temporal analysis of educational homogamy to interpret societal openness, drawing on cumulative insights from mobility research. The validity of the findings from both sides of the exchange is questioned on methodological grounds. Issues raised include the use of inconsistent and incompatible classification categories, aggregation bias, and inadequate temporal design. A trend analysis of assortative mating in Taiwan from 1976 to 1997 further challenges the studies' findings and offers new insights into temporal changes in educational homogamy. The results show significant non-synchronous changes in various components of the educational homogamy parameters. Furthermore, the trend for the intrinsic association parameter has an elongated U-shaped form, indicating support for the status closure argument, rather than an inverted U-curve pattern, a linear trend, or invariance, as has been suggested.

Key words: social mobility, marriage, educational homogamy, comparative research, Taiwan

Research on Societal Openness: From Mobility to Assortative Mating

For students of stratification, the comparative degree of societal openness across countries and its possible temporal trend within individual countries have long been central concerns in empirical investigation (Sorokin 1959). Past research tends to focus exclusively on intergenerational and intragenerational occupational or class mobility (to cite a few, Erikson and Goldthorpe 1993; Featherman et al. 1975; Ganzeboom et al. 1989; Grusky and Hauser 1984; Wong 1990, 1992), as socioeconomic attainment is arguably one of the most important stratification outcomes in modern societies. While the first generation of mobility research examined absolute mobility rates (Lipset and Zetterberg 1959), a second generation research in the 1970s and early 1980s shifted the focus to relative mobility rates using log-linear models (for a brief discussion of the historical development of mobility research, see Ganzeboom et al. 1989; Goodman and Clogg 1992).[1] The working hypothesis shifted accordingly from the thesis of constant absolute mobility to the thesis of constant relative mobility in both cross-national and temporal change studies (Featherman et al. 1975; Hauser, Koffel, Travis, and Dickinson 1975; Hauser, Dickinson, Travis, and Koffel 1975).

Conclusions drawn from this second generation of mobility research are rather profound: aside from dramatic discrepancies in the marginal

[1] Relative mobility rates are measured by the odds-ratios found in the table, which are unaffected by macrostructural changes in the occupational or class distributions.

distributions between the parental and respondent's generations, relative mobility rates are remarkably stable over time and across countries. For convenience, this invariance thesis will be labeled as the *FJH thesis*.[2] Such findings of cross-national and temporal invariance, however, have been challenged since the 1980s by a third generation of mobility research that adopts advanced sophisticated statistical models (for cross-national studies, see Ganzeboom et al. 1989; Wong 1990, 1992; Yamaguchi 1987; for temporal studies, see Hout 1984, 1988; Ganzeboom et al. 1989; Wong 1994; Wong and Hauser 1992; but see Erikson and Goldthorpe 1993 for a different interpretation).[3] Specifically, these studies dispute not only the original formulation of relative invariance in the FJH thesis but also other theses postulating either convergence or universal trends towards increasing or decreasing openness. The cumulative insight seems to be that there are significant cross-national and temporal variations in relative mobility, but they do not follow any simple pattern explainable by certain universal mechanisms and processes. Thus, the role of family socialization, labor market structure, political institutions, and the historical relationship between capital and labor need to be taken into account in examining cross-national and temporal patterns in stratification outcomes.

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[2] The FJH thesis is originally formulated from a cross-national comparison of the mobility patterns found in the United States and Australia (Featherman, Jones, and Hauser 1975) and does not apply to temporal pattern. The temporal invariance component is extended in separate works by Hauser and his associates (Hauser, Koffel, Travis, and Dickinson 1975; Hauser, Dickinson, Travis, and Koffel 1975).

[3] See Wong (1989, 1990) for a thorough critique of the FJH thesis.

Recently, these issues have been revisited by a new line of comparative research on societal openness -- the study of assortative mating or educational homogamy (Smits et al. 1998, 2000; Ultee and Luijkx 1990). According to Smits et al. (1998), research on educational homogamy can illuminate societal openness because it is an *additional* indicator of the same phenomenon illustrated in social mobility research, and because marriage patterns can have important consequences for social inequality. For instance, "[i]f many marriages take place between individuals who differ from each other on socioeconomic characteristics, overall inequality in society may be lower than if few such marriages occur" (p. 265). Thus, the study of marriage homogamy offers a new ground for sociological research on societal openness, and its findings can be used to confirm or invalidate findings from past mobility research.

Smits et al. (1998, 2000) present perhaps the most ambitious investigation of educational homogamy attempted so far. Using data from 60 or more countries, their study generates several interesting findings that challenge conventional wisdom about societal openness. Summarily speaking, Smits et al. (1998) find that there are significant cross-national variations in educational homogamy; the relationship between level of economic development and educational homogamy is non-linear and has an inverted U-shaped form; cultural characteristics are associated with degree of educational homogamy, with Catholic, Muslim, Confucian and mixed Catholic/Protestant countries showing significantly more educational homogamy than Protestant countries; and finally, industrializing societies with a horticultural background exhibit significantly less educational homogamy than industrializing societies with an agrarian background.[4]

Such bold attempt and provocative findings naturally invite controversy and debate. Without directly challenging the general validity of the findings, Raymo and Xie (2000) initiate a critical exchange by making several pointed critiques of the study. First, they question whether and why the four "Confucian" countries analyzed (Hong Kong, Japan, South Korea, and Taiwan) should have a higher level of educational homogamy than other countries. Second, they take issue with the practice of inferring temporal trend from cross-sectional design, arguing in particular that inferring temporal variation from regional variation or "reading history sideways" ignores the possibility of period effects and/or interactions between country and time (Thornton 1992). Finally, they suggest the use of recently married couples, rather than all married couples, in the study of assortative mating because such design provides the ability to "pin down the influence of particular *historical* periods and thus of particular macro-level influences" (Raymo and Xie 2000:774).[5]

Using first marriage tables from four countries (China, Japan, Taiwan, and United States) and at two time points -- one in the early 1970s and the

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[4] The implied inverted U-shaped trend prediction is at odds with past findings from mobility research: the linear trend of increasing openness in a number of industrial societies (Ganzeboom et al. 1989) and the U-shaped trend of restratification in Hungary (Wong and Hauser 1992).

[5] Unfortunately, one of the tables they used (Taiwan in 1975) contains all marriages rather than first marriages (Raymo and Xie 2000). The kind and degree of bias this inconsistency poses to their empirical investigation is unclear.

other at mid- to late 1980s, Raymo and Xie (2000) then present a significantly different pattern of temporal changes in assortative mating within the individual countries and variation across countries.[6] They find that the strength of educational homogamy has decreased in the three "Confucian" countries, but remained stable in the United States. Furthermore, educational homogamy was not particularly higher in the former than the latter in the 1980s.[7] These findings prompt them to conclude that there is some partial support of the FJH thesis that "*relative* openness in all industrialized societies should be roughly the same," and "a general trend toward greater societal openness over time that depends not on the level of economic development but rather one that is characterized by unique cultural paths not easily represented by readily observable characteristics such as dominant religion" (Raymo and Xie 2000:780, emphasis original).

In response, Smits et al. (2000) acknowledge the drawback of their early analysis in "test[ing] the hypotheses about the effect of modernization using

[6] This unbalanced design implicitly assumes that the pattern found in the United States is representative for all western industrial and Protestant societies. In fact, one of the issues raised by Smits et al. (2000:783) is precisely that the United States has a large proportion of Catholics in the population, and is possibly exceptional.

[7] It is possible to further restrict the log-multiplicative layer effect (ϕ) in Raymo and Xie's work and conclude that by the mid- to late 1980s, the level of educational homogamy is the same in the United States, Japan, and China, and significantly lower in Taiwan. This model yields 63 *df* and L^2 of 87.34 (compared to 59 *df* and L^2 of 85.32 in the final preferred model by Raymo and Xie 2000:779). This finding is, of course, at odds with past findings. It is the contention of this paper that such an interpretation is completely unwarranted.

data for countries with different levels of development at a given point in time" (p. 782). But they question that Raymo and Xie's alternative strategy to investigate the problem on "the basis of trend data for four countries is better than testing them on the basis of cross-national data for 65 countries" (p. 783). As a compromise, they conduct additional analysis on 60 of the original 65 countries by dividing the data into two age (older and younger) cohorts. This reanalysis affirms their earlier conclusion that the relationship between educational homogamy and modernization is curvilinear and has an inverted U-shaped form. More particularly, contrary to the findings of Raymo and Xie, no significant change in educational homogamy is observed in any of the four Confucian countries studied originally (note that China is not considered a Confucian society under their classification), and the level of educational homogamy is higher in the younger cohort than the older one in the United States. In short, they find no evidence that the Confucian pattern is converging to the level of Protestant countries. Nor is there any support for the FJH thesis.

Given the advanced statistical techniques and sophisticated modeling both sides adopted, and the openly conflicting and contradictory findings on basic questions of societal openness that have occupied stratification research for decades, this critical exchange is worth examining closely in some details. To put the conflicting results in perspective, I offer in the following section a conceptual and methodological critique of the approach and strategies taken by both groups of researchers, identifying several common drawbacks of their works that would cast serious doubts on the generality and validity of both sets of findings. This is followed by an analysis of educational homogamy in a country studied in both works -- Taiwan -- using a rigorous temporal design. The results will yield a detailed profile of temporal changes, against which the conclusions about temporal trends by the two groups of researchers may be reassessed.

Methodological Problems in the two Comparative Homogamy Research

The Lack of Comparability

A common question in comparative research is whether the results obtained are "real," or, in other words, whether the phenomena being compared are comparable or distinctly different. There are two ways to establish comparability in comparative research: formal equivalence or functional equivalence (Verba 1971). Under formal equivalence, the goal is to strive for the exact measures(s) in each country, which is often difficult because formal equivalence seldom exists in the real social world. An alternative strategy is to establish functional equivalence, which assumes the existence of identical *latent* concepts in various countries, though they may have different empirical indicators for the same concept.[8] The identification of functional equivalence requires theorization and qualitative reasoning (Nießen 1982). Theoretical considerations are necessary for the formulation of relevant general dimensions and their criteria of pertinence. Qualitative

[8] Different terminologies have been used for such a distinction between formal and functional equivalences. Armer (1973), for example, calls them phenomenal identity and conceptual equivalence.

reasoning is required for knowing, identifying, and interpreting context-bound elements (see Garnier and Hout 1976 for an excellent illustration).

In the case of comparative educational homogamy research, to establish functional equivalence, researchers need to provide sound justification for the nature and number of educational categories used in the cross-classified tables.[9] The same reasoning should be applied consistently across countries and should not result in any distortion of the underlying association pattern. Clearly, methodological convenience and the lack of detailed information in the original data source cannot be considered valid ground for arbitrary classification.

Careful readers of the studies by Smits, Ultee, and Lammers, and Raymo and Xie would notice immediately that they adopt dramatically different classifications for different countries, and the categories used for the common

[9] The same also applies to mobility research. But the problem is generally taken care of, as there are standardized classifications and researchers often provide theoretical and empirical justifications for their particular choice. The standardized categories are then applied consistently across countries to facilitate cross-national comparison. For example, some commonly adopted classifications include the neo-Weberian CASMIN class categories by Erikson and Goldthorpe (1993), the neo-Marxian class categories by Erik O. Wright (1985), and the neo-functionalist classification based on skills, socioeconomic status, and/or occupational prestige (Featherman and Hauser 1978; Treiman 1977). In the case of educational assortative mating, it is possible to use a standardized classification such as the expanded CASMIN educational categories in cross-national and temporal comparisons (Müller et al. 1990). This classification, however, is probably more valid for advanced industrial societies than agrarian and industrializing societies.

countries in their analyses also differ significantly (see Table 1 for details). For example, the four categories used by Smits, Ultee, and Lammers for China are no education, primary, junior middle, and senior middle or higher; and for Japan, primary or less, junior high school, senior high school, and junior college or higher. Raymo and Xie, on the other hand, use primary or less, junior high school, senior high school, and university for China; and junior high school, senior high school, junior college, and university for Japan. Similar discrepancies are also noticeable for Taiwan. It is obvious that such differences in classifications will not only affect the degree of educational homogamy found within individual countries, but also make cross-national comparison and cross-validation futile and misleading.

Table 1. Inconsistent Educational Categories Used in Cross-National Comparisons

	Smits, Ultee, and Lammers	Raymo and Xie
(a) China	No Education	Primary
	Primary	Junior High School
	Junior Middle	Senior High School
	Senior Middle or Higher	University
(b) Japan	Primary or Less	Junior High School
	Junior HS	Senior High School
	Senior HS	Junior College
	Junior College or Higher	University
(c) Taiwan	< 4 Years	Primary
	4-6 Years	Junior High School
	7-12 Years	Senior High School
	> 12 Years	University
(d) United States	< 12 Years	< 12 Years
	12 Years	12 Years
	13-16 Years	13-15 Years
	17+ Years	16 Years

If the choice of classification is critical, then it is important to examine the reasoning and justification behind the particular choices of the two groups of researchers. By treating education as a positional good, Smits et al. (1998:271-2) argue that each society can have its own classification to reflect the types of distinctions that are important in that particular context. Although Raymo and Xie offer no explicit defense of their choice of classification, it is reasonable to assume that they implicitly adopt the same *relative* education position, especially given their study's orientation as a critical extension of the former. Interestingly, however, they do not give any justification for their different choices of categories. This omission makes it impossible to assess the validity of their categories versus those used by Smits, Ultee, and Lammers.

The discrepancies in the categories used illustrate a vital weakness in the researchers' practical use of the relative education classification strategy, even though it is defensible in theory. In actual practice, unless the criteria used in specifying the classification are clearly reasoned and formulated, inconsistent and erratic choices by different researchers may distort the relative standing of educational attainments within any particular society, leading to systematic biases in the association between husbands' and wives' education. Given the different categories used, this is probably the case in either one or both of the studies. However, since the severity of the biases introduced is unknown, it is unclear whether they can account for the inconsistent findings. In short, any conclusions about an inverted U-curve relationship between educational homogamy and industrialization, or other relationships with political institutions and dominant religion (Smits et al. 1998, 2000), or a pattern of cross-national invariance at high levels of industrialization (Raymo and Xie) cannot be established unless the country-specific categories have been demonstrated to be indeed "comparable." To be fair, Smits et al. (1998:272) try to assure readers that their relative educational categories do not introduce any biases or significantly alter the results of their findings. However, in the discussion below, I will illustrate that their strategy in ascertaining this is not foolproof. In sum, arbitrary and inconsistent classification has rendered the two studies' findings about comparative educational homogamy questionable and unreliable.

Aggregation Bias

In their critical comment, Raymo and Xie (2000) raise concern against Smits, Ultee, and Lammers' finding of an inverted U-shaped relationship between educational homogamy and economic development partly by citing the well-established evidence of a post-World War II trend of increasing educational homogamy in the United States (Kalmijn 1991; Mare 1991). It is rather striking then that Raymo and Xie (2000) themselves arrive at a finding of invariance that equally contradicts the well-established evidence, especially when their data are derived from Mare (1991). What methodological lesson can we draw from this peculiarity? A careful examination of the works reveal two major differences in data treatment: (a) while Mare (1991) uses only raw counts (that is, no standardization procedure), the sample size has been standardized in Raymo and Xie's study; and (b) the original 5 by 5 first marriage tables have been aggregated to form 4 by 4 tables in Raymo and Xie's work. Either one or both of these differences could have contributed to the different findings. In the following discussion, only the second difference will be examined in details, as it points to a common problem in the studies of Raymo and Xie, and Smits, Ultee, and Lammers.[10]

Except for maintaining some kind of comparability with the work of Smits, Ultee, and Lammers, there is little reason for Raymo and Xie to restrict their analysis to the highly aggregated 4 by 4 classifications. The decision to combine categories definitely requires careful deliberation, since an artificial restriction of analysis to a limited number of highly aggregated categories may alter the association in significant ways. This problem is formally known as aggregation bias (see Goodman 1981 for a detailed discussion). In general, when categories are combined in any cross-classified tables, there should be a *minimal* loss in association between row and column variables. If the loss is large, the aggregation bias can distort results in significant ways and even the adoption of statistically powerful 1-*df* tests cannot salvage the situation.

To investigate the extent of aggregation biases in the work of Raymo and Xie, and Smits, Ultee, and Lammers, we can examine the consequence of such aggregation. Table 2 reports the loss of association from the original 5 by 5 tables in Mare (1991) to Raymo and Xie's aggregated 4 by 4 tables. Note that it is the two lowest educational categories (less than 10 and 10-11 years of schooling) that are combined. The two columns under panel A

[10] Evaluation of the former practice is also more difficult because there is no conventional standard for treating unequal sample sizes in tabular data analysis.

report the log-likelihood ratio test statistics for the independence model. In terms of absolute loss, all are statistically significant at 0.05 level or less. In terms of relative loss, it is 20.8% in 1940, 10.8% in 1960, 0.05% in 1970, 3.8% in 1980, and 3.0% in 1985-87. While the degree of relative loss is small in the last three tables (1970 and later), it is significant in the first two periods.

Table 2. Aggregation Bias in the Analysis of First Marriages in the United States

		Log-Likelihood Chi-Square Statistic (L2)			
Year	N	5 x 5	4 x 4	5 x 5	4 x 4
		(A) Independence Model		(B) Quasi-Independence Model	
		16 df	9 df	11 df	5 df
1940	4051	1919.00	1519.77	433.23	201.12
1960	8934	4226.95	3768.39	1348.09	915.25
1970	13153	6343.63	6027.45	2018.56	1519.19
1980	13154	6597.80	6345.18	1839.23	1381.05
1985-87	3957	1946.38	1888.96	497.06	396.61

Note: The five educational categories are <10, 10-11, 12, 13-15, and 16+ and the four categories are <12, 12, 13-15, and 16+.

Since the extent of homogamy is usually high in most societies, researchers often include diagonal parameters to capture this tendency. The last two columns under panel B are models of quasi-independence and provide additional information about possible loss in association. The comparison reveals that not only is the absolute loss statistically significant at all times, the amount in relative loss is also considerable (53.6%, 32.1%, 24.7%, 24.9%, 20.2%, respectively). Again, the extent of relative loss is much more apparent in earlier than later periods. This result indicates that a substantial degree of heterogamy (that is, marriage of persons with dissimilar characteristics) may be mistaken as homogamy in the aggregated table. In other words, when diagonal parameters are included in statistical modeling, such parameterization would make the American society look more open or have a higher degree of educational homogamy in the highly aggregated tables than the original tables. The differential loss of association at various times would further prevent any meaningful interpretation of temporal trends and raise serious doubts about any conclusion based on the highly aggregated 4 by 4 tables.

To explore further how aggregation bias may affect the study of educational homogamy in countries that have more refined classification of educational categories, Table 3 reports the same exercise on the first marriage tables of Taiwan from 1976 to 1997 (Ministry of Interior, Republic of China, various years). Unlike Raymo and Xie (2000), the analysis here does not include the 1975 table because it covers all marriages, not just first marriages. This strategy is consistent with their argument that analysis of newlyweds are more preferable in trend analysis of assortative mating (see also Kalmijn 1994; Mare 1991; Qian 1997). The original classification has six categories: college or more, senior high school, junior high school, primary, self-taught, and illiterate.[11] Raymo and Xie (2000) combine the last three categories to form the highly aggregated 4 by 4 tables. For comparison, an intermediate 5 by 5 table that combines only the self-taught and illiterate categories is also included here. Because the number of marriages is large in each year, the cell

[11] In 1976, the self-taught category is replaced by the literate category, but they should be referring to the same group of individuals.

counts have all been adjusted downward by a factor of 100. This should have no impact on assessing the relative loss in association.

Table 3. Aggregation Bias in the Analysis of First Marriages in Taiwan

Year	N	6 x 6	5 x 5	4 x 4	6 x 6	5 x 5	4 x 4
		(A) Independence			(B) Quasi_Independence		
		25 df	16 df	9 df	19 df	11 df	5 df
1976	1466	629.78	626.72	604.45	293.12	285.14	135.64
1977	1502	661.75	659.03	643.45	293.34	286.69	160.62
1978	1569	672.38	669.78	656.69	294.26	288.09	181.23
1979	1485	599.03	596.05	587.68	263.31	258.24	183.00
1980	1639	619.59	615.33	603.88	289.10	279.45	210.09
1981	1640	565.80	561.54	551.55	271.53	264.96	214.56
1982	1516	490.28	487.11	470.47	235.07	229.92	191.20
1983	1508	464.35	457.41	442.75	220.02	210.29	173.11
1984	1423	423.53	415.20	401.61	195.77	183.29	152.66
1985	1436	398.36	391.97	379.20	178.61	178.55	144.84
1986	1356	353.10	344.19	328.91	161.17	149.42	123.34
1987	1347	318.27	310.41	297.32	141.81	131.74	105.74
1988	1450	286.01	280.61	269.68	119.03	113.45	94.18
1989	1472	279.45	275.34	264.92	112.92	108.41	89.84
1990	1319	237.31	232.35	223.08	89.80	84.84	66.24
1991	1516	245.43	241.14	233.00	92.84	88.95	69.00
1992	1579	269.13	264.26	255.17	74.70	70.15	55.44
1993	1417	303.86	297.60	288.96	82.20	75.80	53.19
1994	1569	384.21	374.28	365.74	95.89	83.85	48.95
1995	1475	330.13	317.65	310.32	89.71	77.51	40.88
1996	1529	338.05	323.44	316.49	84.75	73.53	32.55
1997	1536	391.90	376.76	366.45	95.74	84.09	38.99

Note: The six categories are college+, senior high school, junior high school, primary, self-taught, and illiterate, the five categories combine self-taught and illiterate whereas the four categories combine primary, self-taught, and illiterate.

From panel A of Table 3, it is clear that combining the original 6 by 6 tables into the two other aggregated tables resulted in limited loss of information. If we treat the rescaled sample sizes as real counts and the difference in the log-likelihood chi-square statistics distributes like a chi-square distribution, then none of the loss in association between bride's and groom's education is statistically significant at the 0.05 level. The same conclusion, however, does not apply to the quasi-independence models under panel B. While the loss of association from the 6 by 6 to the 5 by 5 table is negligible, both absolute and relative loss in the highly aggregated 4 by 4 tables cannot be ignored. More significantly, the loss is not evenly distributed, high in early and later years but moderate in the period between. The proportional loss of association ranges from a low of 18.7 percent in 1982 to about 60 percent in 1996 and 1997.

Since the statistical models adopted by both groups of researchers in analyzing educational homogamy include diagonal parameters, the implication drawn from Table 3 is highly significant. It suggests that the use of highly aggregated 4 by 4 tables in Raymo and Xie's analysis may have significantly reduced the extent of association in Taiwan. If we assume that the loss of association in their 1975 table follows the pattern shown here, then there would be over 50 percent loss of association, while the loss is only about 26 percent in 1990. This means that their study would have artificially created a much more open image of the Taiwanese society in the mid-1970s. A substantial amount of association between those without formal education (self-taught and illiterate) and those with only primary education would have been mistakenly treated as homogamy rather than heterogamy or off-diagonal association. In sum, our exercise here clearly illustrates the danger of significant *over-estimation* of societal openness in certain periods when highly aggregated tables are used with diagonal marriage homogamy parameters, as in Raymo and Xie's study. This casts serious doubts on the validity of any evidence about temporal changes found therein.

The same problem probably applies to the analysis by Smits, Ultee, and Lammers as well, since they report having more refined educational categories in a number of countries studied. However, without access to the original data, it is difficult to empirically examine the extent of possible aggregation bias in their works. Although Smits et al. (1998:272 footnote) reassure readers that the extent of aggregation bias is minimal because the correlation between the Spearman rank order correlations from the highly aggregated tables and the original tables is exceptionally high ($r = 0.98$), this assurance is on rather shaky ground, since we have already demonstrated that the loss in association does not necessarily lie only in the overall association but can be in specific locations as well.

Table 4. Spearman Rank Order Correlations Under Different Aggregation of First Marriages in Taiwan, 1976-97

	No Cells Deleted			Diagonal Cells Deleted		
Year	6x6	5x5	4x4	6x6	5x5	4x4
1976	0.565	0.565	0.582	0.353	0.330	0.047
1977	0.576	0.576	0.587	0.311	0.294	0.056
1978	0.572	0.572	0.579	0.281	0.266	0.066
1979	0.554	0.554	0.560	0.248	0.237	0.071
1980	0.538	0.537	0.542	0.243	0.230	0.086
1981	0.513	0.513	0.517	0.218	0.210	0.093
1982	0.499	0.499	0.502	0.205	0.197	0.094
1983	0.477	0.477	0.483	0.180	0.171	0.078
1984	0.467	0.468	0.472	0.168	0.157	0.069
1985	0.453	0.454	0.459	0.145	0.138	0.064
1986	0.433	0.434	0.439	0.132	0.122	0.045
1987	0.413	0.413	0.419	0.115	0.104	0.020
1988	0.376	0.377	0.382	0.071	0.064	-0.013
1989	0.367	0.368	0.373	0.064	0.056	-0.028
1990	0.352	0.352	0.358	0.048	0.040	-0.055
1991	0.337	0.338	0.342	0.044	0.035	-0.067
1992	0.335	0.335	0.337	0.003	-0.004	-0.057
1993	0.378	0.379	0.381	0.050	0.037	-0.031
1994	0.400	0.401	0.402	0.059	0.043	-0.039
1995	0.374	0.374	0.377	0.083	0.063	-0.038
1996	0.359	0.359	0.362	0.071	0.053	-0.049
1997	0.398	0.399	0.401	0.082	0.066	-0.031

Pearson Correlation Between Spearman Rank Order Correlations

	No Cells Deleted			Diagonal Cells Deleted		
	6x6	5x5	4x4	6x6	5x5	4x4
6 x 6	1.000			1.000		
5 x 5	1.000	1.000		0.999	1.000	
4 x 4	1.000	1.000	1.000	0.820	0.838	1.000

To further illustrate the inadequacy of their treatment, Table 4 calculates the Spearman rank order correlations for the Taiwanese data from 1976 to 1997 under both the independence and quasi-independence models. For the full tables with no cells blocked or fitted exactly, the Spearman rank order correlations remain virtually the same under various levels of aggregation. In fact, the Spearman correlations under the highly aggregated tables are even slightly higher than the original 6 by 6 tables. As shown at the bottom of Table 4, the Pearson correlations between various Spearman rank order correlations are virtually equivalent. However, when we examine the columns with the diagonal cells deleted or fitted exactly, the story is very different. While the Spearman rank order correlations are virtually the same under the 6 by 6 and 5 by 5 tables, the same is not true of the highly aggregated 4 by 4 tables. The latter correlations are more similar only in recent years, but the disparities are large in the early period, especially between 1976-1987. The relatively poor results for the highly aggregated 4 by 4 tables can be further illustrated in their low correlations with the two less aggregated tables. The correlations are in the order of 0.82 whereas the correlations between the less aggregated tables are very close to 1.00.

To recapitulate, our analysis demonstrates that there are serious aggregation biases with the highly aggregated tables used by Raymo and Xie, leading to significant loss of association found in the original tables. The loss is perhaps even more significant in the analyses by Smits, Ultee, and Lammers, given that they report the existence of more refined categories in a number of countries studied. In any case, even if there is little loss in the overall level of association, notable loss in specific locations would still lead to significant biases in the findings.

The Illusion of Temporal Trends

The strongest critique Raymo and Xie (2000) raise against Smits et al. (1998) is that their use of cross-national variations to infer historical tendencies may have masked real temporal trends. How can temporal trends be properly studied is indeed an important issue that is all too easily sidestepped in studies dealing with data from multiple countries, as Raymo and Xie inadvertently did in their own study. Analyzing only two time points in each country, they lay claim to "true trend data, with two marriage cohorts per country, while SU&L had only cross-sectional data" (Raymo and Xie 2000:773-774, emphasis original). Interestingly, Smits et al. (2000) follow suit in their reply, assuming that having two age cohorts is adequate to study temporal trends. An attentive analyst will realize that having only two time points in an analysis hardly permit anyone to infer temporal trends, not to say that the time span is 10 to 15 years apart. The infinite variations that are possible within the lengthy period between, as well as before and after the two time points can no way be seen from such a design. In other words, while Raymo and Xie correctly criticize Smits, Ultee, and Lammers for the problem of "reading history sideways," they have substituted it with an equally problematic strategy of compressing history to linearity in their study. So have Smits, Ultee, and Lammers in their reply. Important information about temporal trends within individual countries may thus have been missed by both studies.

While Smits, Ultee, and Lammers' argument for the inclusion of more countries in comparative study to minimize the risk of distortion owing to country specificities is well taken, the danger of compounding distortions with questionable extrapolation of temporal trends in individual countries should also be noted. Before any meaningful cross-national comparison can be made, we need to make sure that the trends observed in individual countries are real rather than illusory, and can stand up to rigorous scrutiny.

Trend of Assortative Mating in Taiwan: Change or No Change?

To see whether the “trends” found by the two groups of researchers hold true in a rigorous temporal study of educational homogamy, annual data of first marriages in Taiwan, a country included in both studies, are analyzed here for the period 1976-1997.[12] With 22 observational time points, this analysis gives a detailed profile of temporal changes that should enable us to assess whether and to what extent the level of educational homogamy in Taiwan remained largely unchanged or decreased, as the findings of Smits et al. (2000) and Raymo and Xie indicate respectively. Since our earlier test on aggregation bias shows that the 5 by 5 tables result in little loss in association, and the combination of the two categories of self-taught and illiterate into a

[12] The ideal data set for studying the relation between economic development and educational homogamy in Taiwan would include first marriages in the 1960s and early 1970s, the “takeoff” period of rapid economic growth in the country. The extended period of study would have also meant a closer match with the time frame in Smits et al. (2000) analysis, which “runs roughly from somewhere in the 1940s to somewhere in the 1970s” (p.786). Unfortunately, there are no available data on first marriages in Taiwan before 1976. Nonetheless, the findings here should still allow us to observe if there is any sign of the inverted U-curve patten described by Smits, Ultee, and Lammers.

broader category of no formal education is reasonable for a society with universal primary education, these tables are used in the analysis.

Table 5. Models for Educational Homogamy in Taiwan, 1976-97

Model Description	*df*	L^2	*BIC*	*ID*	*p*
1. Conditional Independence	252	9118.16	5458.46	19.95	0.000
2. Conditional Quasi-Independence	242	3598.46	1082.42	9.63	0.000
3. Full Two-Way Interaction	336	503.79	-2898.56	3.64	0.000
4. Log-Linear Layer Effects	315	214.20	-3060.82	2.34	1.000
5. Log-Multiplicative Layer Effects	315	309.64	-2965.38	2.80	0.575
6. Log-Linear Layer Effects Model with Linear Trend Restriction	335	274.75	-3208.21	2.73	0.993
7. Log-Linear Layer Effects Model with Non-Linear Trend Restriction	334	225.54	-3247.02	2.44	1.000
8. Heterogeneous RC and DIAG	88	68.51	-846.42	0.52	0.939
9. Homogeneous RC and DIAG	340	546.70	-2988.24	3.64	0.000
10. Simple Heterogeneous RC and Heterogeneous DIAG	214	228.36	-1996.58	1.73	0.239
11. Simple Heterogeneous Equal RC and Heterogeneous DIAG	217	259.37	-1996.76	1.88	0.026
12. Non-Linear ϕ & DIAG4, Linear DIAG1, DIAG2, DIAG3, & DIAG5, Homogeneous μ_i and v_j	332	267.17	-3184.60	2.40	0.996
13. Non-Linear ϕ & DIAG4, Linear DIAG1, DIAG2, DIAG3, & DIAG5, Homogeneous μ_i and v_j (with equality restriction in trend coefficients for DIAG1 & DIAG2, and DIAG3 & DIAG4 but with sign changes)	334	267.54	-3205.02	2.42	0.997

Note: See text for explanation.

Table 5 presents a series of statistical models applied to test for temporal changes.[13] The first model (line 1) is the conventional baseline model for

[13] All models are estimated by the LEM program written by Vermunt (1997), and the input and output files are available from the author. As in Table 2, all cell counts have been

comparison. It postulates conditional independence over time, and, as expected, this null association model does not fit the data. The model of conditional quasi-independence (line 2) postulates that aside from a tendency of educational homogamy, marriage between people with dissimilar educational backgrounds occurs randomly. Although the model does not fit well, it indicates that close to 60 percent of the association can be explained by similar spousal characteristics, leaving about 40 percent to association with dissimilar characteristics. Thus, the result indicates that *both* homogamy and heterogamy are important to our understanding of educational assortative mating in Taiwan.

The model of no temporal change in assortative mating (line 3) fits the data moderately well, and can explain close to 94 percent of the association between spousal characteristics (336 df, L^2=504). This is the conventional test for the FJH temporal invariance thesis. Although the fit of the model looks acceptable, this model cannot be accepted as evidence for invariance without the formulation of an alternative test (Wong 1989, 1990).

Model 3 postulates that the complete set of odds-ratios is constant across all time points, that is,

$$\theta_{ijk} = \theta_{ij} \tag{1}$$

where θ_{ij} represents the adjacent odds-ratios for row i and column j, and θ_{ijk}

divided by 100. Thus, the log-likelihood ratio chi-square statistics cannot be interpreted as the conventional chi-square statistic. The choice of various models is based largely on the *BIC* statistic, the Bayesian information criterion (Raftery 1995). In general, models with the more negative *BIC* values are preferred over others.

represents the adjacent odds-ratios for row i, column j, and layer k. Under equation (1), the complete set of odds-ratios does not depend on the layer variable (year) and is constant over time.

Two 1-*df* tests have been developed to test whether the complete set of odds-ratios is indeed identical across tables. The first one is the log-linear layer effects model (LL1), which postulates that the odds-ratios (θ_{ijk}) have the following relationship:

$$\theta_{ijk} = \theta_{ij}\ \beta_k \tag{2}$$

where β_k represents the log-linear layer effects with the normalization that $\beta_1 = 0$ (Yamaguchi 1987; Wong 1990). The contrast between equations (1) and (2) offers a k-1 *df* test. On the other hand, the log-multiplicative layer effect model (LL2) postulates that the odds-ratios have the following relationship:

$$\log\ \theta_{ijk} = \psi_k\ \log\ \theta_{ij} \tag{3}$$

where ψ_k represents the log-multiplicative layer effects, subject to normalization that the sum of the squared of ψ_k equals to 1 (Xie 1992; Powers and Xie 2000). The contrast between equations (1) and (3) also offers a k-1 *df* test. The major difference between LL1 and LL2 lies in the specification of the layer effects. Though LL2 has the advantage that any interchange in the row and/or column categories will have no effect on the fit of the model, it does not follow that LL2 always provides a better fit and is therefore preferable to LL1.

Lines 4 and 5 present results from LL1 and LL2, respectively. Although both models offer significant improvement over their invariance counterpart (line 3), the log-linear specification (L^2=214) clearly offers a better fit than the log-multiplicative alternative (L^2=310). An inspection of the log-linear layer effects coefficients indicates a visible trend. Two alternative tests are used to capture this trend: a linear restriction (line 6) and a non-linear (quadratic) restriction (line 7). In terms of the changes in goodness-of-fit statistic and *BIC*, they both offer significant improvement over the unrestricted model, but the one with nonlinear trend restriction offers a more parsimonious and appropriate interpretation.

The close fit between the observed estimates from the unrestricted model and the fitted estimates from the non-linear trend model can be seen in Figure 1. According to the parameter estimates reported in the top panel of Table 6, the complete set of odds-ratios does not seem to follow Smits, Ultee, and Lammers' "hypothesized" trend of an initial increase in the early period and a gradual decline later. Even if we allow for the possibility that the hypothesized trend would spread over a period longer than that studied here, a match with the observed trend is still questionable. For instead of a gradual decline, there was a rapid decline in the odds-ratios from the mid-1970s to the late 1980s. Since then, the odds-ratios have become stabilized and begun to increase gradually after 1991. In fact, if the trend maintains its course after 1997, the odds-ratios in year 2000 can be extrapolated to be as high as it was in 1986.

For more information on the pattern of temporal changes, it is important to look beyond whether the complete set of odds-ratios has changed in the same order over time and investigate how each specific components may have changed. Sometimes, it is possible for individual components to not only experience different rates of change over time, but also offset each

Table 6. Selected Parameter Estimates of Educational Homogamy in Taiwan

	Base	Linear Trend	Quadratic Trend
A. Global Odds-Ratios:			
Log-Linear		-0.0391	0.0012
Layer Effects		(0.0037)	(0.0002)
B. Specific Parameters:			
Homogamy			
College +	2.1080	-0.0295	
	(0.1710)	(0.0065)	
Senior HS	-0.7718	0.0295	
	(0.0654)	(0.0065)	
Junior HS	0.3902	0.0135	
	(0.0713)	(0.0049)	
Primary	-0.7104	-0.0135	0.0030
	(0.1378)	(0.0049)	(0.0004)
Other	1.4989	-0.0580	
	(0.1479)	(0.0121)	
Intrinsic	4.3478	-0.2073	0.0042
Association (ϕ)	(0.1469)	(0.0280)	(0.0014)
Estimated Row and Column Scores:			
	Bride	Groom	
College +	-0.5437	-0.5771	
	(0.0229)	(0.0218)	
Senior HS	-0.4799	-0.4587	
	(0.0261)	(0.0260)	
Junior HS	0.0632	0.1345	
	(0.0139)	(0.0135)	
Primary	0.5468	0.5776	
	(0.0218)	(0.0208)	
Other	0.4136	0.3238	
	(0.0214)	(0.0232)	

Note: The standard errors for the specific parameters in the lower panel are obtained by the jackknife method. All coefficients are statistically significant at the 0.01 level or less.
See text for details.

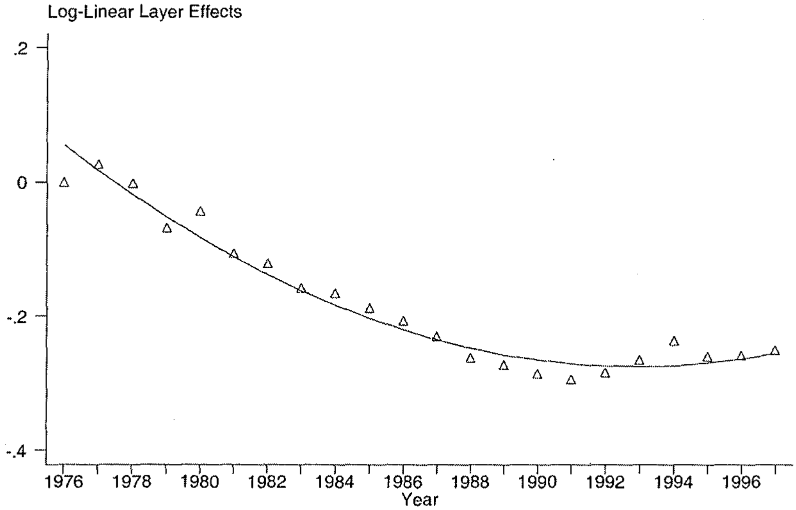

Figure 1. Observed and Fitted Log-Linear Layer Effects

other's changes, a result that may not be detected in the global tests. Since we have already established the importance of both diagonal and off-diagonal association in understanding assortative mating in Taiwan, our statistical model should try to separate the two and examine their changes over time.

The log-multiplicative row and column effects model with diagonal-specific effects is chosen here to represent specific components of educational homogamy. The model has the following form:

$$\log m_{ijk} = u + u_i + u_j + u_k + u_{ik} + u_{jk} + \phi_k \mu_{ik} v_{jk} + \delta_{ijk} \qquad (4)$$

where m_{ijk} is the expected frequency for cell (i,j,k), u_i, u_j, u_k, u_{ik}, and u_{jk} are marginal parameters, subject to normalization, ϕ_k represents the intrinsic association between wife's and husband's education, μ_{ik} represents the estimated distance between wife's educational categories (row scores), v_{jk} represents the estimate distance between husband's educational categories (column scores), and δ_{ijk} represents the diagonal-specific educational homogamy parameters. Note that equation 4 postulates that all assortative mating parameters vary over time. Our goal is to impose various restrictions to these parameters, including invariance, linear trend, and quadratic trend, in order to find credible evidence about possible changes over time.

Equation 4 represents the model in line 8 and its fit is satisfactory. Its homogeneous counterpart (line 9), on the other hand, shows a significant deterioration of fit. The contrast between the two confirms our earlier finding that at least some of the assortative mating parameters have changed over time. To facilitate our search for temporal trends, the model in line 10 postulates that only the intrinsic association parameters (ϕ) vary over time but not the row and column score parameters. The overall fit of this model is satisfactory and the deterioration of fit is relatively minor. Note that model 10 assumes asymmetrical association. Its symmetrical counterpart, model 11, which imposes equal row and column scores, results in significant deterioration of fit. Thus, our finding is consistent with that of Smits et al. (1998) about the existence of asymmetrical association between wife's and husband's education.

To avoid lengthy presentation of the search for changes in individual parameters, only two final models are presented.[14] The model in line 12 postulates non-linear trends for the intrinsic association parameter (ϕ) (Wong

[14] Alternative specification that permits uniform change in all diagonal assortative mating parameters has been tested as well. But the fit of the model is not as good as the one presented here. The result is available from the author on request.

1995) and the diagonal marriage homogamy for primary education (DIAG4), and linear trends for the remaining four educational categories (DIAG1, DIAG2, DIAG3, and DIAG5). Model 13 further equates the trend coefficients of the diagonal marriage homogamy for college or more (DIAG1) and senior high school (DIAG2), and for junior high school (DIAG3) and primary education (DIAG4), except with sign changes. Both models have a satisfactory fit. The deterioration in fit of model 13 over model 12 is marginal, and the improvement in the *BIC* statistic is notable. Therefore, model 13 is chosen to represent temporal changes in assortative mating in Taiwan.

The values of the estimated parameters and their corresponding jackknifed standard errors under model 13 are presented in panel B of Table 6. All coefficients are statistically significant at 0.01 level or less. The result clearly indicates that not all measures of the assortative mating parameters move in the same direction. While the degree of marriage homogamy declined gradually for those with college or more education and those without formal education, there is simultaneously a visible trend of increasing marriage homgamy among those whose educational attainment are somewhere between the two extremes, that is, those with (senior and junior) high school and primary education. In fact, among those who received only primary education, the rate of increase has accelerated rapidly over time. That is, as general education became more prevalent and widespread for all Taiwanese residents, only those who are better educated (college or more) experienced greater latitude in choosing their marital partners whose education can be quite different from their own. For the less educated population, the tendency to marry alike actually increased over time, particularly among those who received only primary education (see Figures 2 to 7 for a visual display of their changes over time).

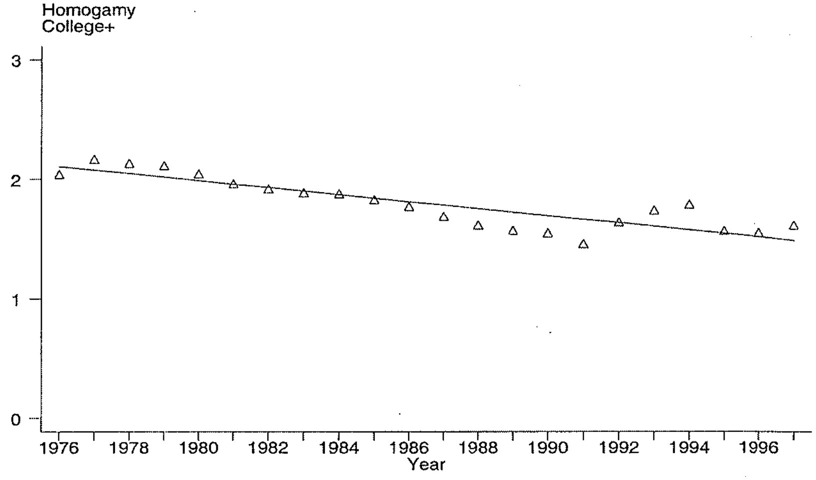

Figure 2. Observed and Fitted Homogamy--College or More

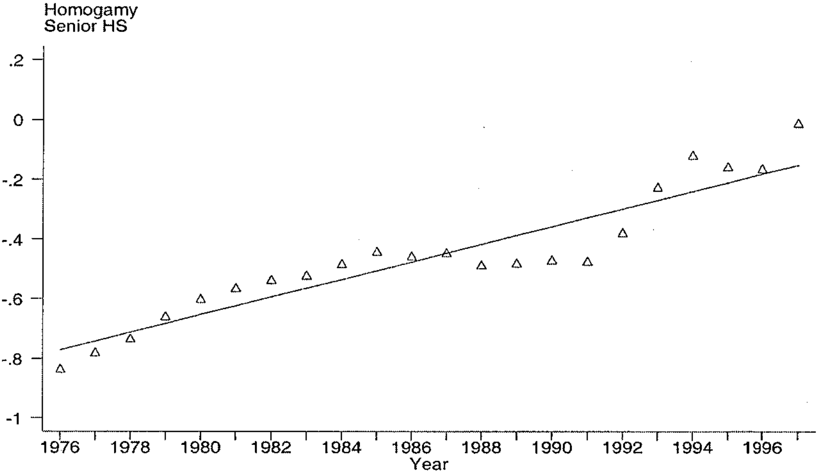

Figure 3. Observed and Fitted Homogamy--Senior HS

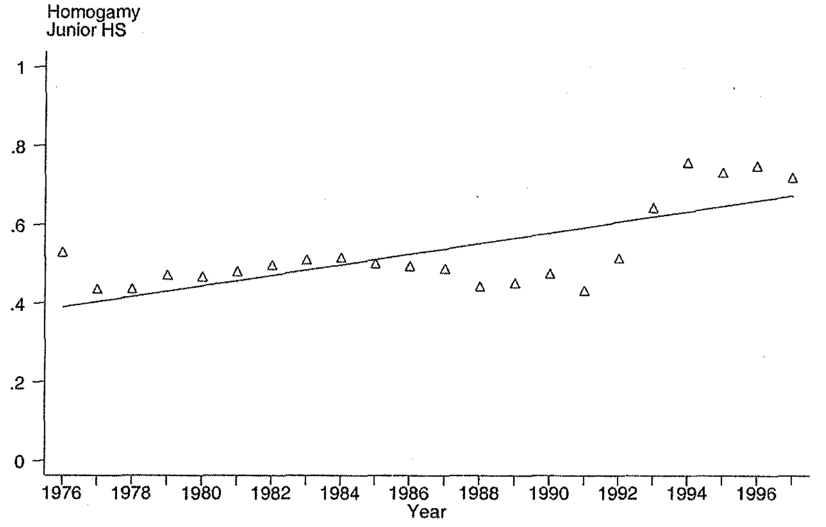

Figure 4. Observed and Fitted Homogamy--Junior HS

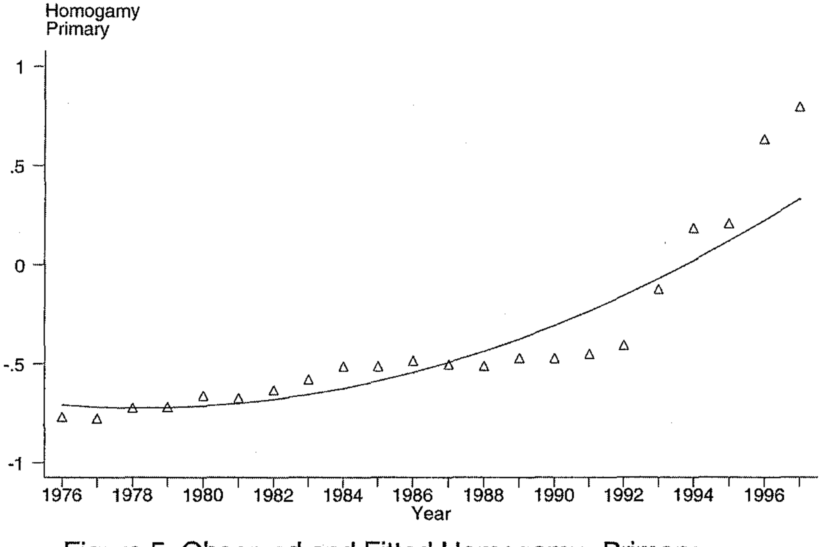

Figure 5. Observed and Fitted Homogamy--Primary

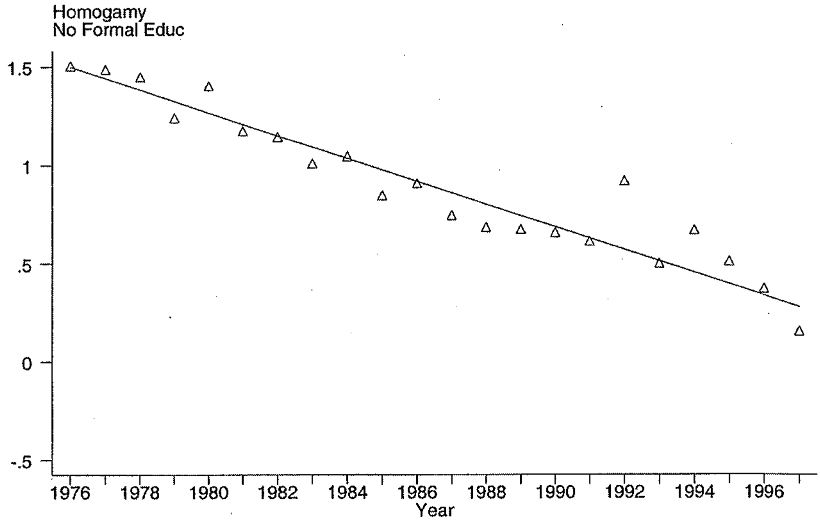

Figure 6. Observed and Fitted Homogamy--No Formal Educ

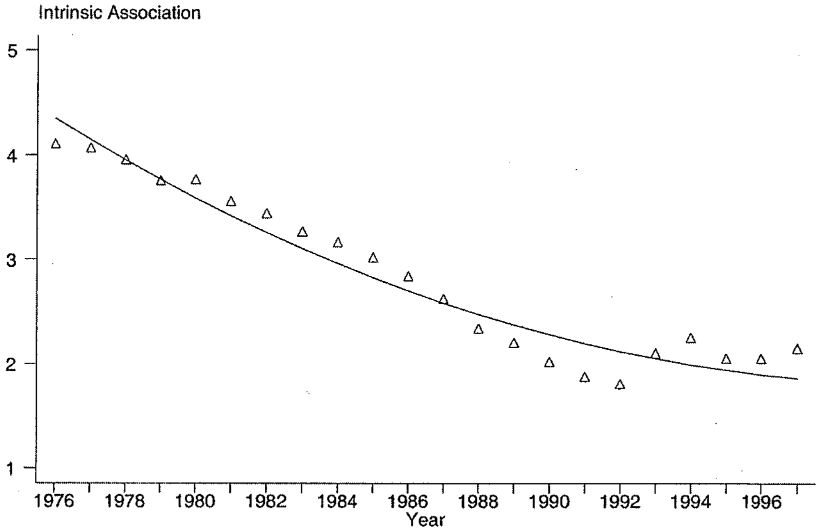

Figure 7. Observed and Fitted Intrinsic Association

On the other hand, the trend for the intrinsic association parameter (ϕ) displays a rather different pattern. The intrinsic association parameter indicates a dramatic and steady decline over time but the rate of decline is decelerating. Again, this pattern does not seem to conform to the inverted U-shape pattern found in Smits et al. (1998). The trend of decline may look similar to the finding of Raymo and Xie (2000), but it reveals a decelerating rate of decline that is unobservable in their two-time-points study. If we interpret the size of intrinsic association parameter as an indicator of societal openness, it is clear that after a long, sustained period of increasing openness, contemporary Taiwan is showing increasing signs of resistance to further openness. In fact, should the trend maintain its course after 1997, the intrinsic association parameter will reach its lowest level in year 2000 and start to increase thereafter, a pattern that is consistent with the post-World War II American experience (Mare 1991; Kalmijn 1991).

Conclusion

The temporal analysis of assortative mating in Taiwan in this study bears out the cumulative wisdom from mobility research that temporal variations in societal openness do not follow a simple pattern attributable to some universal mechanisms or processes such as modernization. Indeed, the findings make clear the poverty and reductiveness of any formulation about universal pattern -- be it invariance, increasing openness, decreasing openness, or trendless fluctuation (Sorokin 1959) -- and its supposed relationship with macrostructural and cultural factors. The non-linear (elongated U-shaped) trend found here may or may not apply to other countries, but it documents temporal changes that challenge Smits, Ultee, and Lammers' conclusion about an inverted U-shaped relationship between economic development and educational homogamy. It also casts serious doubts on Raymo and Xie's conclusion about partial support for the FJH thesis. The study thus rings a cautionary bell against the blanket use of broadly defined economic and cultural/religious factors to cover over and explain away significant cross-national differences and temporal changes in educational homogamy. If the interaction between economic development and Confucianism is to account for the level of educational homogamy in Taiwan or other countries, as both the studies of Smits, Ultee, and Lammers, and Raymo and Xie claim, then it is necessary to first hypothesize how and under what kind of mechanisms Confucianism influences individual choice of marital partners based on educational background, then operationalize the hypothesis and test it explicitly. Without such explicit hypothesis and testing, any conclusions about the relationship between Confucianism and the level of educational homogamy are bound to be more obfuscating than illuminating.

The importance of looking beyond universal patterns to attend to specific processes and mechanisms is further evident from the finding here that various components of the educational homogamy parameters show different patterns of temporal change. While some of the parameters have increased in strength, others have decreased. Furthermore, their rates of change are not necessarily constant over time. Such non-synchronous change makes clear the danger of losing important information and of potential bias in using single summary measure to study assortative mating and draw conclusions about cross-national patterns and temporal trends in societal openness.

Finally, the trend of assortative mating in Taiwan found here parallels some previous findings on intergenerational mobility in Hungary (Wong and Hauser 1992). Specifically, the results are consistent with predictions from the status closure argument (Parkin 1971) rather than the thesis of industrialism or modernization. In other words, the findings support the argument that while modernization and industrialization weaken past established social structures and relationships, and promote greater freedom for individuals to pursue personal goals and choice of marital partners, there are also significant limits to these changes and the consequent increase in societal openness. As new stratification and status systems develop and strengthen, the restrictions on individual choices will increase and consolidate accordingly. What these specific changes are and whether they will be strong enough to reverse the current trend depend on many factors. Further insights into these changes can be gained from more research on a selective set of theoretically interesting countries over an extended period of time, rather than large-scale cross-sectional comparative projects that try to include as many countries as possible.

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Appendix A. Frequencies of First Marriages in Taiwan

	College+	Sr HS	Jr HS	Primary	Other	College+	Sr HS	Jr HS	Primary	Other
			1976					**1982**		
College+	5923	2359	400	281	41	7365	3621	725	350	212
Sr HS	6956	11598	3296	2875	291	10686	19336	7737	4336	806
Jr HS	1956	6692	4795	7500	738	2961	13612	15893	11116	1476
Primary	1708	10328	12065	44167	4723	1150	7749	14805	16690	2501
Other	247	1411	2092	11301	2868	317	1317	2299	3453	1076
			1977					**1983**		
College+	7315	8382	2230	1564	217	7829	3899	784	382	419
Sr HS	2722	13839	7950	10173	1182	10783	20040	8847	4291	1231
Jr HS	463	3948	5742	12003	1872	3002	13917	17197	9936	1688
Primary	292	3674	9398	40684	8993	1089	7107	14386	13623	2267
Other	53	342	922	4233	2034	442	1332	2177	3066	1075
			1978					**1984**		
College+	8216	9842	2458	1663	226	7564	3981	851	399	389
Sr HS	3131	16126	9549	9944	1180	9933	19568	9696	3997	1273
Jr HS	525	4858	7332	12468	1933	2908	13203	17828	8766	1643
Primary	332	4252	11103	36655	7548	940	6224	13431	11128	1929
Other	79	504	1086	4069	1819	403	1285	2193	2641	1060
			1979					**1985**		
College+	7968	2990	550	341	102	8091	4437	1056	413	468
Sr HS	9984	16510	5230	4390	593	10199	20788	10948	4052	1314
Jr HS	2482	9939	8758	11328	1178	2946	13110	18653	8363	1566
Primary	1416	8947	12636	29196	3454	939	5761	12960	9266	1730
Other	240	1180	1913	5839	1362	447	1208	2086	2057	777
			1980					**1986**		
College+	7972	11242	2914	1455	332	7281	4082	1033	396	485
Sr HS	3420	19308	12362	9517	1352	9474	18903	11059	3716	1416
Jr HS	588	6469	12099	14867	2399	2895	12272	18366	7651	1538
Primary	332	4608	12645	26973	5482	938	5389	12041	8018	1512
Other	155	686	1419	3653	1676	479	1361	2190	2192	933
			1981					**1987**		
College+	7889	3611	630	362	197	7233	4300	1222	453	534
Sr HS	11572	20077	7265	4613	848	9226	19069	11484	3857	1540
Jr HS	3172	13743	14770	12560	1629	2881	12063	18091	7234	1632
Primary	1357	9103	15946	21810	3013	1033	5213	11120	7027	1653
Other	342	1336	2392	4436	1278	607	1502	2417	2288	979

Appendix A (Continued). Frequencies of First Marriages in Taiwan

	College+	Sr HS	Jr HS	Primary	Other	College+	Sr HS	Jr HS	Primary	Other
			1988					**1993**		
College+	7084	4619	1413	517	540	9777	4958	2072	464	436
Sr HS	9428	20585	13417	4604	1750	9147	21169	13772	2226	951
Jr HS	3179	12998	19516	7508	1802	4387	16268	28167	4598	882
Primary	1265	5848	11930	7141	1726	984	3203	6257	2527	463
Other	670	1748	2646	2063	969	934	2331	3302	1926	544
			1989					**1994**		
College+	7201	4669	1481	628	604	12189	5507	2446	364	472
Sr HS	9532	20245	13398	4661	1824	10518	23794	15172	1844	760
Jr HS	3084	12974	19371	7562	1977	5386	17944	33607	3717	752
Primary	1513	6158	11853	7466	1809	935	2527	4790	2074	323
Other	741	2040	2925	2353	1149	1171	2963	4462	2506	683
			1990					**1995**		
College+	6361	4099	1400	601	514	9983	5024	2789	396	516
Sr HS	7959	17333	11631	4229	1741	8131	17354	12790	1699	835
Jr HS	2858	11196	16990	6654	1668	5467	16326	34321	3882	1060
Primary	1522	6205	10824	7062	1913	1016	2604	5289	2522	506
Other	816	2017	2758	2353	1162	1378	3553	5816	3221	981
			1991					**1996**		
College+	6785	4718	1638	727	616	9797	4566	3244	302	508
Sr HS	8842	19793	13027	4794	1969	7358	15045	12770	1068	748
Jr HS	3379	13334	18763	7167	2029	6585	18355	43773	2588	1031
Primary	1986	7574	12464	8067	2354	632	1691	3401	1656	417
Other	952	2599	3602	2892	1479	1442	4064	7831	3055	948
			1992					**1997**		
College+	8981	10161	4714	1590	651	14312	6036	3863	293	601
Sr HS	5082	23783	18965	5032	1395	8856	19198	12885	936	620
Jr HS	2100	15767	28775	8678	1863	6596	16446	34185	1667	832
Primary	662	3716	6397	3691	1233	569	1225	2260	1186	250
Other	528	1303	1340	856	598	1781	5403	9533	3234	868

Note: The row and column variables refer to bride's and grooms education, respectively.

To See or Not To See: Another Look at Research on Temporal Trends and Cross-National Differences in Educational Homogamy

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中文摘要

本文透過既有社會流動研究累積的見解，檢視最近一個以教育同質婚配的跨時與跨國分析用於衡量社會開放性的學術辯論。由於辯論雙方在方法論上均存在若干問題，本研究質疑他們對教育同質婚配分析的效度。其中較有疑義之議題包括雙方在教育分類上並不一致，也不相容；教育類別歸屬上的偏誤，以及跨時分析在研究設計上不夠適切。本研究分析台灣從1976至1997年婚配同質性的趨勢，分析的結果進一步挑戰上述雙方的研究發現，同時也對教育同質婚配的變遷研究提供新的洞察視角。資料分析的結果顯示，教育同質婚配模型的各個參數間，存在著非同步性的變異。而且，在教育同質婚配模型中，內在關聯參數的變化呈現一條平滑、細長的U字曲線，意味著地位封閉之論點獲得資料的支持，而不是相關研究所認為的——呈現一個倒U字曲線；一個線性趨勢；或是一個不變模式的各種變化形態。

關鍵詞：社會流動、婚姻、教育同質婚配、比較研究、台灣