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Andras Gabos, Robert I. Gal* and Gabor Kezdi:

Fertility Effects of the Pension System and Other

Intergenerational Transfers:

Test on Hungarian Data

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Introduction

There is a discrepancy between the consumption path and income path of the life-cycle.¹ The elderly as well as the children have to consume, income, however, is produced in the active period. All societies exploit the occurrence of overlapping generations and use two-way flows of transfers, one from the active to children and another one from the active to the elderly, in order to smooth out the difference. The chain of transfers persists due to reputation effects (Kotlikoff, Persson and Svensson 1988) or a reward-or-punish strategy (Hammond 1975). In both models transfers are only backward flowing, from the active to the elderly, that is, $t+1$ period pensions are paid on condition of contributions to t period pensions. However, in both models fertility and forward flowing transfers, such as family allowances, are exogenous. The self-enforcing intergenerational constitutions by Cigno (1993) or Rangel (2003) make $t+1$ period pensions depend also on t period transfers flowing to children rendering fertility choices and the choice over forward flowing transfers endogenous. The chain of intergenerational transfers sustains if transfer flows of the opposite direction are connected.

In a traditional society the institution organizing this chain is usually the extended family. In modern societies such transfers flow among social generations rather than family generations. This historical shift creates a larger risk pool (a comparative efficiency of the family and the insurance market see in Kotlikoff and Spivak 1981), makes intergenerational transfers more easily enforceable and offers insurance against unintended infertility (Sinn 2004). However, if alternative vehicles of wealth accumulation, such as the capital market or social security, offer higher yields some generations may be tempted to desert from the family chain leaving their parents without old-age income and decreasing their fertility (Cigno 1993). As a matter of fact, a full-fledged capital market could, in principle, increase fertility by offering a chance for specialization in accumulating either in physical capital or human capital, much like comparative advantages induce growth in the ricardian theory of international trade (Razin and Sadka, 1995). Yet, this potential can be exploited only if costs and yields are internalized in household utility functions not allowing intra-generational redistribution from the fertile to the infertile. If such positive externalities of raising children can be demonstrated, the negative effect on fertility still persists.

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Bental (1989) consider the internalization by making pension benefits linked to the number of children raised. Abío, Mahieu and Patxot (2003) make declining fertility endogenous by distinguishing between female and male labor and having female labor more complementary to capital than male labor. Capital accumulation makes female wages grow faster than household income rendering the opportunity cost of child raising ever higher. A benefit formula that relates pensions to the number of children functions as a corrective tax and restores the optimal rate of population growth. Kolmar (1997) differentiates between IPAYG systems, where $t+1$ period pensions are paid on condition of contributions to t period pensions, and CPAYG systems, where $t+1$ period pensions are paid according to the number of children raised. In an IPAYG a backward flowing transfer is linked to a previous backward flowing transfer. In a CPAYG a backward flowing transfer depends on a previous forward flowing transfer, much the same way as in the traditional family chain. Kolmar demonstrates the optimality of corner solutions: either pure IPAYG or pure CPAYG are preferable. Van Groezen, Leers and Meijdam (2003) combine forward and backward flows of transfers in order to internalize the positive externalities of child raising. Pensions of phase $t+1$ depend on phase $t+1$ contributions. The discounted value of the latter is given to parents in phase t as child allowance. Fenge and Meier (2003) compare the alternatives of extending the PAYG pension system by a fertility-related component to the introduction of family allowances. Following an Aaronian argument they find that the choice depends on the relation of the targeted fertility rate to the interest rate. If the former is lower family allowances are preferable.

The empirical evidence on the fertility effects of public intergenerational transfers, such as pensions and family benefits, are generally supportive. Usually, however, pensions and family benefits are tested separately. On the fertility effects of pensions Nugent (1985) and Nelissen and van den Akker (1988) offer detailed reviews. Most tests use aggregate data in cross section (e.g. Hohm, 1975 compares 67 countries, Nugent and Gillaspay, 1983 34 Mexican counties, Entwisle and Winegarden, 1984 48 and 52 middle and low income countries in two separate models). A notable exception is Jensen (1990), who, testing Cain's "lexicographic safety first" model on data of the Rand Malaysian Family Life Survey, uses household data. He finds that couples using contraceptives had significantly higher life insurance and expected external, non-family source of income for old age. Cigno and Rosati (1996) and Cigno, Casolaro and Rosati (2003) use macro data but in a time series, not in cross section. They explain savings and fertility by social security coverage in Italy, Germany, the UK and the US. Child benefits appear as control variables in their models.

As for the fertility effects of family benefits, Gauthier and Hatzius (1997) review the empirical literature. These analyses are based on Becker's theory of demand for children. In this field time series analysis on aggregate data is more frequent. Ermisch (1988) used aggregate time series data for Britain, Whittington, Alm and Peters (1990) for the United States, Zhang, Quan and Meerbergen (1994) for Canada, while Gábos (2003) for Hungary. Gauthier and Hatzius (1997) also analyzed time series aggregate data but for 22 countries. Blau and Robins (1989) and Whittington (1992) use a household panel, while Milligan (2002) examines a micro-database from the Canadian Census. The common finding of the literature is a positive but weak effect of family benefits on fertility. The positive relationship is present in all of these analyses, while magnitudes of these effects vary across countries. All these models, however, focus only on forward flowing intergenerational transfers.

The paper is organized as follows. In Section 1 we introduce the data and variables of the model. In Section 2 the model is specified. Finally, in Section 3 we show the results of an alternative way of normalization of the explanatory variables. In particular, employing a standard decomposition technique we extract components of the usually applied explanatory variables that are relevant for individual fertility choices and clean them from components that are inappropriate for the analysis.

1. Trends in fertility, pensions and family benefits in Hungary

We measure fertility with the total fertility rate. The *TFR* is one of the most commonly used indicators of fertility and is defined as sum of age-specific birth rates calculated for the reproductive period of women:

$$TFR = \sum_{i=15}^{49} f_i, \text{ and } f_i = \frac{B_i}{P_i^f}, \text{ where}$$

f_i – age-specific birth rate of women at age i ,

B_i – number of births given by women at age i ,

P_i^f – number of female population at age i .

Total fertility rate for a given year thus expresses the number of children a women would have if she went through her reproductive ages with age-specific birth rates characterizing that calendar year. It must be highlighted however, that *TFR* does not represent the fertility behav-

ior of a real cohort, but of a hypothetical one, since it describes the patterns of one particular year. This index filters out the effects of changes in gender and age composition on the number of births. Using the total fertility rate as a dependent variable in similar models is wide spread in international literature.

Pension and family benefit expenditures relative to the GDP are used as explanatory variables in the model. They operationalize intergenerational transfers flowing forward and backward among overlapping generations. Family benefits can be listed in the first group, while pensions belong to the second. Almost all studies, cross-sectional or longitudinal they may be, apply their relative size as explanatory variables. Nevertheless, one can argue that the pension rate and the family benefit rate depend on some factors that may and some others that may not have an impact on fertility. We show an exercise of a potential decomposition of the explanatory variables and test the result in an alternative regression model in Section 3.

Diagram 1
Pensions/GDP and the total fertility rate, 1950-2002

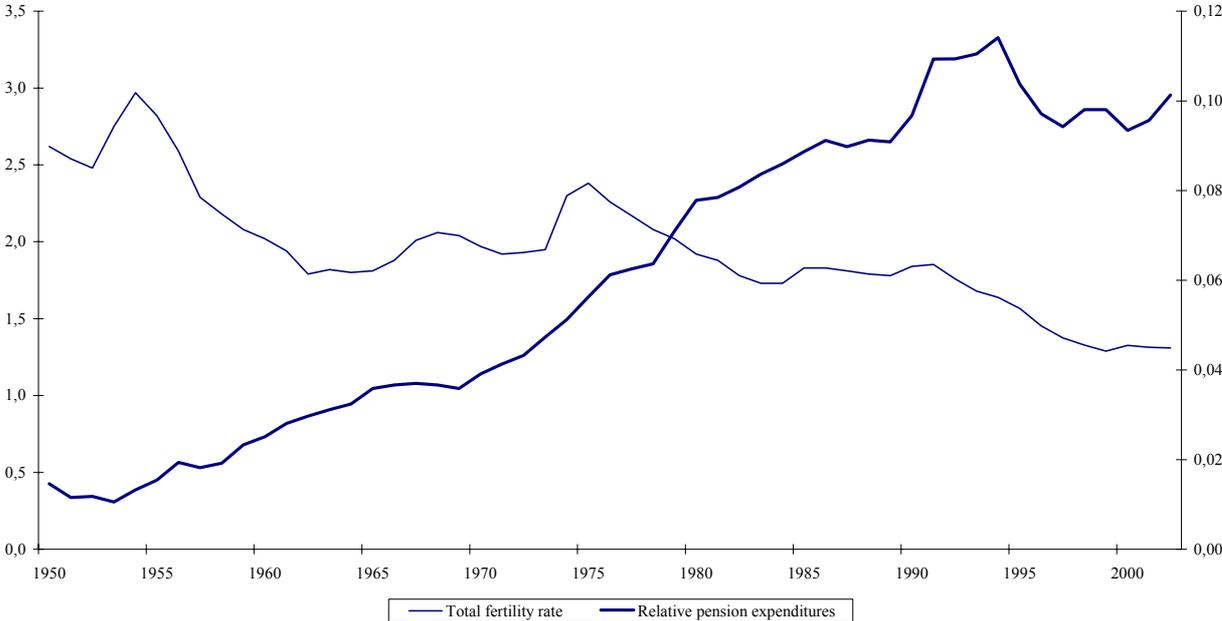


Diagram 1 shows the development of pensions and fertility. The decline of Hungarian fertility in the second half of the 20th century has been fundamentally consistent although we have noticed temporary trends that vary from this both in calendar years and in terms of completed fertility indices.

The funded pension scheme collapsed in World War II and in the inflation that followed. It was re-established in 1950 as a PAYG scheme. The pension rate basically stagnated from the re-start until the extension of eligibility in 1957 and it is only from then that stable growth began. The annual growth rate increased following the 1970 implementation of annual pension indexation. By the beginning of the 1980s the system has become mature and the growth of the pension rate slowed down. From the end of the 1980s, the beginning of the mass influx of those under retirement age, until the mid 1990s it again began to grow significantly. In 1995-1996 the pension rate fell sharply; since then it stagnates.

At first glance the diagram suggests that the rare and small decreases of the pension rate do not increase fertility. In contrast, however, frequently, though not always, significant jumps coincide with drops in the total fertility rate. The other important observation is that fluctuation of the pension rate can be less linked directly to specific institutional changes. We should also add that without filtering out effects of the control variables the above visual examinations would only be suitable for providing some kind of first impression.

Diagram 2
Family benefits/GDP and the total fertility rate, 1950-2002

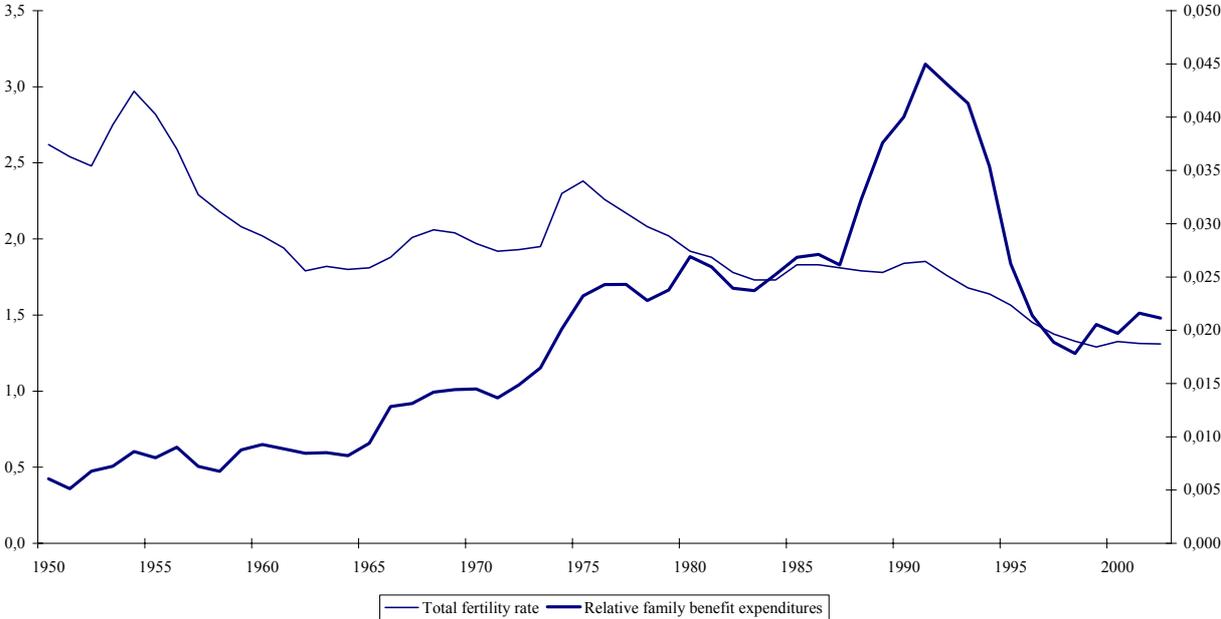


Diagram 2 shows parallel trends in fertility and the family benefit rate. For visual observation the relation seems to be stronger. The growth segments of the fertility series, five in number are linked to the important changes in the family benefit system. This was so despite the fact that the two trends were in opposing directions, that is, we noticed a negative correlation between fertility and spending on family related transfers throughout the entire period.

From among the periods of growth in question the first, that is, 1953-1955 coincides with the complete ban on abortions. The second stage, 1965-1969, contains the implementation of an extended (up to three years) childcare allowance in 1967. Stage three (1973-1975) coincided with the 1973 implementation of a complex population policy program, which contained yet another tightening of abortion practices. The package contained positive incentives such as significant housing support for couples with children and increasing the real value of cash family benefits. Stage four (1984-1985), although lasted only one year, more or less coincided with the introduction of yet another new form of support, the wages related child care fee in 1985. Compared to the previous stage, however, there is a significant difference. The growth period is not followed by an instant and relatively quick decrease, but rather by stagnation. Finally the last growth stage, 1988-1990, was preceded by a substantial increase of cash family related transfers, above all, child allowances.

The intuitive comparison of growth periods of fertility and family policy indicates that the existence of family benefits and the continuous changes thereof affected fertility. However, we cannot determine whether the effect was temporary or lasting, did it merely affect the timing of births or did it have effect on completed fertility. The relationship appears obvious, however, that negative changes caused significant breaks in the fertility trend with effects short lived and influencing only the timing of births. Indeed, completed fertility for the most affected cohorts in the 1953 ban on abortions was lower than that of younger cohorts despite the very high calendar year fertility index between 1953-1955. In contrast, measures via positive incentives suggest longer lasting changes in trends.

Despite the various explanations, all fertility theories agree that the decrease in fertility in industrial societies is related to the labor market participation of women that is the growth of the opportunity cost of child raising. This can be approached with several variables such as female wages, the level of women's education or the rate of economic activity of women. Among these, data availability allows to use economic activities of women for the period since 1950. Gender based time series data on wages is missing. For female employment we used a variable that generated the ratio of the number of women over the age of 25 who were economically active to the total population of women in their fertile age.

Employment, and within this female employment, steadily rose until around the late 1980s in Hungary. In the early 1990s employment decreased significantly for several years and since the end of the 1990s stagnation and very slow growth have been noticeable. Consequently, the female employment trend is not linear and so is its relation with fertility. This problem was dealt with by including the square of the employment variable as well. Since female employment indices for the 1950s were not available (with the exception of 1950) we linearly intrapolated the data for this period, that is, we assumed that the value changed the same amount every year between 1950 and 1960. Since this process artificially puts information in a data series, distorted results are possible. In the interest of determining the magnitude of this effect we ran alternative models as well without female employment variables.

Infant mortality is used in nearly all studies as a control variable. The effect of a high infant mortality rate can be twofold: more live births are necessary to reach the desired number of children while high mortality can detour parents' desire to rear children. The first effect seems to be much stronger, so we expect the relation between infant mortality and fertility to be positive. When specifying the variable we assumed that infant mortality depends on the amount spent on child healthcare on one hand and the income of households with children on the other.

The relationship between income and child rearing is a crucial question to the economic theory of fertility. These theoretical predictions are not unambiguous pertaining to the direction of the relationship. Income effect is positive, however, the substitution effect results in a negative relation between income and fertility. In addition, infant mortality and household income are highly correlated. Thus, we used only one of them, infant mortality, in the model.

We also took the number of marriages per 1000 persons into consideration. There are arguments both for and against using the marriage rate as a control variable. Marriage is not a prerequisite to having children, however the majority of births still take place in wedlock.² The Hungarian regulation, although in ever decreasing areas, still prefers marriage to cohabitation. Thus, married couples are at an advantage compared to those in cohabitation, for example when applying for housing support, citizenship or adoption. The analysis of changes in fertility cannot be separated from changes in the marriage rate. We expect that the relationship

² This is still so despite the fact that the ratio of children born out-of-wedlock is on the rise. Out-of-wed-lock fertility is lower than that of married couples.

between marriage and fertility is positive: increases in the marriage rate appear in the increases to the fertility index and vice versa.

As suggested marriage is not a prerequisite for child rearing, so the decrease in the marriage rate in western societies may appear not to be a direct cause of decreased fertility, rather it is the manifestation of a common cause behind both processes. This mutual cause is particularly complex, perhaps the easiest way to understand it is via the changes that have taken place in the social position of women. When assuming such a common reason we also have to take into account that the existence of family benefits, in addition to the direct effects on fertility, also influence decisions pertaining to marriage and thus indirectly decisions to rear children. This assumption however questions the necessity of using the marriage rate as control variable. In as much as our assumptions are correct the influence of family benefits on fertility are stronger in those models, which do not contain marriage rates.

Fertility behavior also depends on supply via costs of avoiding pregnancy. In the following we have incorporated this effect in the model in a very simplified form. We strove to eliminate the shock-like changes caused by the stringent abortion policies of the early fifties and early seventies when developing the annual fertility rate. Thus, we have included a binary variable (ABORT) in the models. This produced a value of 1 between 1953 and 1955 and between 1973 and 1975, and 0 in all other cases. We are aware that this variable is not capable of capturing the costs of avoiding pregnancy. The appropriate data (e.g. the spread of birth control pills and other contraceptives) however, are not available.

2. Econometric model and estimation results

We would like to measure the effects of intergenerational cash transfers on fertility. That is, we are interested in fertility responses to exogenous changes in transfers. The data at hand are national time-series from Hungary. Given the non-experimental nature of our data, we need to worry about endogeneity issues and try to handle them. With the time series at hand, we have to be even more careful in order to avoid estimating spurious effects because of nonstationarity and with stationary series, to take autocorrelation into account. In this section, we outline the estimation strategy and the main results.

According to Figure 1, log total fertility rate ($\ln TFR$) is trending. Augmented Dickey-Fuller and Phillips-Perron tests cannot reject the null of unit root, indicating a stochastic trend (around a linear trend). At the same time, however, the KPSS test cannot reject the null of trend-stationarity. Log pension rate ($\ln P$) and log family benefits ($\ln FB$) each contain a unit

root, and here all tests are in accord.³ We shall run all of our models on first differenced variables, but among the robustness checks, we also run everything on detrended $\ln TFR$ as a dependent variable.

In our preferred models, the first difference of the log total fertility rate ($\Delta \ln TFR$) is regressed on its own lag(s), the lagged first difference of the rate of family benefits to GDP ($\Delta \ln FB$) and lagged first difference of the rate of pension expenditures to GDP ($\Delta \ln P$) – and their further lags. Equation (1) shows the simplest such model. As year t fertility is a result of decisions made at least 9 months earlier, only year $t-1$ or earlier family benefit changes or pension rate changes can have a causal effect.

$$(1) \quad \Delta \ln TFR_t = \alpha_0 + \sum_{s=1}^S \alpha_s \Delta \ln TFR_{t-s} + \sum_{s=1}^S \beta_s \Delta \ln FB_{t-s} + \sum_{s=1}^S \gamma_s \Delta \ln P_{t-s} + u_t$$

In our estimated models, only first lags are going to be significant. In practice (1) therefore simplifies to

$$(1a) \quad \Delta \ln TFR_t = \alpha_0 + \alpha_1 \Delta \ln TFR_{t-1} + \beta_1 \Delta \ln FB_{t-1} + \gamma_1 \Delta \ln P_{t-1} + u_t$$

The estimation results show that we need at most one lag of the left-hand side variable. Estimated long-run effects of FB and P are, therefore, the following:

$$(2) \quad LE(FB) = \frac{\sum_{s=1}^S \beta_s}{1 - \alpha_1}, \quad LE(P) = \frac{\sum_{s=1}^S \gamma_s}{1 - \alpha_1}$$

Or, if only first lags are significant, the long-run effects are:

$$(2a) \quad LE(FB) = \frac{\beta_1}{1 - \alpha_1} \quad LE(P) = \frac{\gamma_1}{1 - \alpha_1}$$

Standard errors of the long-run effects can be estimated via the delta-method approximation:

³ No cointegration of $\ln TFR$ with either $\ln FB$ or $\ln P$ (or both).

$$(3) \quad V\left[g(\hat{\boldsymbol{\theta}})\right] \approx \nabla g(\boldsymbol{\theta})'_{\boldsymbol{\theta}=\hat{\boldsymbol{\theta}}} V(\hat{\boldsymbol{\theta}}) \nabla g(\boldsymbol{\theta})_{\boldsymbol{\theta}=\hat{\boldsymbol{\theta}}},$$

where $\hat{\boldsymbol{\theta}}$ is the vector of estimated coefficients, g is the function as defined in (2) or (2a), ∇g its gradient (evaluated at the point estimates), and $V(\hat{\boldsymbol{\theta}})$ is the variance-covariance matrix of the estimated coefficients.

Our major problem is whether we can give a causal interpretation to the estimated coefficients and the long-run effects calculated from them. If changes to family benefits or the pension rate are exogenous to unmeasured innovations to fertility a year later, then the causal interpretation is appropriate. If either one (or both) is endogenous the estimated effects cannot be interpreted in a causal way. We have no good instruments for either potentially endogenous right-hand side variable, but we follow some other strategies to get at the problem.

First, we use year t FB and P variables as controls for unobserved changes in fertility that might have occurred simultaneously with benefit and pension changes:

$$(4) \quad \Delta \ln TFR_t = \alpha_0 + \alpha_1 \Delta \ln TFR_{t-1} + \beta_0 \Delta \ln FB_t + \beta_1 \Delta \ln FB_{t-1} + \gamma_0 \Delta \ln P_t + \gamma_1 \Delta \ln P_{t-1} + v_t$$

Since only lagged benefit and pension changes can have causal effects current changes are proxies for correlated unobservables. Therefore, the lagged effect estimates (β_1, β_2 , etc, γ_1, γ_2 , etc.) and the long-run effect estimates based on them are cleared of the endogeneity (provided current effects are perfect proxies). Even if they are imperfect proxies, including them in the equation reduces the potential bias from endogeneity.

Second, we use a set of potentially relevant control variables. This strategy is useful if at least some of the endogeneity in family benefit and pension rate changes can be captured by proxies for omitted variables. The included controls are factors that affect fertility on their own rights, and therefore all correspond to year $t-1$. Moreover, since all controls are nonstationary themselves, they are entered in differences as well.

$$(5) \quad \Delta \ln TFR_t = \alpha_0 + \alpha_1 \Delta \ln TFR_{t-1} + \beta_1 \Delta \ln FB_{t-1} + \gamma_1 \Delta \ln P_{t-1} + \boldsymbol{\delta}' \Delta \mathbf{x}_{t-1} + w_t$$

The controls include female employment rate, its squared, infant mortality, marriage rate, and a dummy variable for the 1953-1955 and the 1973-1975 period, when strict abortion rules were introduced after and before a more liberal regime. These variables are thought to

have substantial effect on fertility, both on theoretical grounds and according to the previous empirical literature as we briefly reviewed in the introduction.

Table 1. Aggregate time-series estimates of the effect of family benefits over GDP (*FB*) and pensions over GDP (*P*) on the total fertility rate (*TFR*) next year. All variables enter the regression in differences, *TFR*, *FB* and *P*, crude marriage rate and infant mortality in log differences

Dependent variable: $\Delta \ln TFR_t$	Equation (1)	Equation (4)	Equation (5)
$\Delta \ln FB_{t-1}$	0.237 (0.057)	0.185 (0.061)	0.245 (0.050)
$\Delta \ln P_{t-1}$	-0.304 (0.079)	<u>-0.225</u> (0.087)	-0.347 (0.069)
$\Delta \ln TFR_{t-1}$	0.180 (0.127)	0.177 (0.129)	-0.119 (0.125)
Constant	Yes ^a	Yes ^a	Yes ^a
Current $\Delta \ln FB_t$ and $\Delta \ln P_t$	No	Yes	No
Other controls	No	No	Yes
Number of observations	52	52	52
R ²	0.45	0.51	0.64
Serial Correlation LM test (χ^2) (p-value)	0.16 (0.92)	0.58 (0.75)	3.88 (0.14)
Long-run effect of FB	0.289 (0.101)	0.369 (0.116)	0.219 (0.060)
Long-run effect of P	-0.371 (0.084)	-0.389 (0.102)	-0.310 (0.086)

Note: Regression coefficients are **significant at the 1% level**, significant at the 5% level and *significant at the 10% level*.

^a Variable entered but not significant at the 5% level.

The time series of the sample run from 1950 to 2003. After differencing and taking into account that lags are used for the estimation, we are left with 52 observations. Given the relatively short time-series, the differenced nature of the series, and the relatively large number of parameters (especially in (4) and (5)), any significant estimates indicate strong correla-

tions. Small degrees of freedom are also likely to result in unstable results across specifications unless the underlying relationships are robust – or, in other words, finding robust estimates are a remarkable result given the relatively short time series. Table 1 shows the main results.

The estimates of the family benefits (FB) are significant at the 1% level in all specifications, and so are the estimations of the pension rate (P) except for specification (4) where it is significant at the 5% level. As we shall discuss later, our preferred results are the estimates from specification (5). According to the point estimates, *1 percent increase in family benefits is associated with a 0.25 percent increase in total fertility*. The pension estimate indicates that *1 percent increase in the pension rate is associated with a 0.35 percent decrease in total fertility*.

The model gives a relatively strong fit for a regression on first differences. The dynamic specification is also correct, as the tests indicate no remaining serial correlation after entering the once-lagged dependent variable. No further lags of the right-hand-side variables are significant either in any of the specifications. In fact, even the first lag of the TFR is not significant at the 5% level. We include it anyway in order to have comparable results with later specifications, in which one lag is usually significant.

Estimated effects of pensions are higher (in absolute terms) when proxies are controlled for. Adding contemporary FB and P changes to the regression (equation 4) leads to reduction, supporting our interpretation that they control for otherwise unobserved heterogeneity.

In order to evaluate endogeneity of changes in family benefits and the pension rate, one has to understand the political economy of those changes. Indeed, increasing family benefits in certain periods are reactions to decreasing trends in fertility, known from government publications and supported by a negative correlation between changes in family benefits and lagged changes in fertility. In a similar vein, growing fertility in a fixed institutional setting also results in higher public spending on child benefits. Provided changes in fertility are positively correlated (which they are), simultaneity introduces a downward bias into our baseline estimates on family benefits. Although possible fertility effects of the pension system are not directly considered in policy decisions, we cannot rule out a reverse relationship between the pension rate changes and fertility.

The timing of the events allows us to assess the relative importance of the two directions in the simultaneous relationship between intergenerational transfers and fertility. Granger-causality tests can tell us whether lagged family benefits or pensions help explaining

current fertility, and whether lagged fertility can help us explaining current transfers. Joint evidence for the first (i.e. that transfers Granger-cause fertility) and against the second (i.e. that fertility does not Granger-cause transfers) would imply that the effects estimated by our regressions are dominated by the causation from transfers to fertility. Table 2 shows the results of pairwise Granger-causality tests run on log differences:

Table 2: Granger-causality tests between fertility and family benefits over GDP, and fertility and pensions over GDP

Null Hypothesis:	Obs	F-Statistic	Probability
$\Delta \ln \text{FB}$ does not Granger Cause $\Delta \ln \text{TFR}$	51	3.229	0.049
$\Delta \ln \text{TFR}$ does not Granger Cause $\Delta \ln \text{FB}$		1.445	0.246
$\Delta \ln \text{P}$ does not Granger Cause $\Delta \ln \text{TFR}$	51	1.753	0.185
$\Delta \ln \text{TFR}$ does not Granger Cause $\Delta \ln \text{P}$		2.215	0.121

The tests provide evidence that family benefits Granger-cause fertility but not the other way round. This supports the claim that most of the cross-correlations between fertility and family benefits is coming largely from the effect of the latter on the former. The implications for the pension rate are more mixed: we cannot establish Granger-causality in either direction. Besides showing that fertility does not affect pension rates much, the results indicate that causality from the pension rate to fertility is also less robust than from family benefits.

The results indicate a robust and most likely causal relationship between family benefits and fertility. The magnitude of the effect is moderate, with an elasticity around 0.20-0.25. The pension rate also seems to affect fertility, by a magnitude of 0.20-0.35. The causal interpretation of the latter association is, however, less certain.

3. Normalizing family benefits and pensions in a different way

The pension system decreases the fertility motives if more people in their reproductive period have pensioners in their reference group (coverage), and these pensioners are more independent of support from their children (replacement). In case of low coverage the decision maker is not influenced by the experiences of pensioners in her environment, and in case of low replacement rate the strategy to prepare for old age outside the family is not very attractive. The

underlying decision model does not order the decision maker to be able to imagine herself in unprecedented institutional circumstances, such as the introduction of a PAYG pension scheme, but only as having the life of the current elderly in her later years. In a stationary world this behavior is rational. In a non-stationary world this way of operationalization is less demanding of the decision maker.

This implies an explanatory variable of the model different from the one usually applied. Most models, as our baseline model above, use the rate of pension expenditures to the GDP. However, this indicator may blur factors affecting the fertility choice with others, which may not. The relative size of pension expenditures can be broken down into five components as follows:

$$(6) \quad (Bb/\text{GDP}) = (B/O)(O/A)(b/w)(A/E)(w/(\text{GDP}/E)),$$

where B is the number of pensioners, b is the average pension, Bb is the total pension budget, O is the number of people above retirement age, B/O is coverage, A is the number of people in active age, O/A is old-age dependency, w is average wage, b/w is the replacement rate, E is the number of contributors, A/E is the reciprocal of activity rate and $w/(\text{GDP}/E)$ is the reciprocal of wage efficiency.⁴

Out of these five components only two, coverage and replacement, affect directly the fertility choice. The other three components refer to incidences which either may not affect fertility decisions or for which the micro-level effects are not obvious. The activity rate blurs female and male employment even though these two factors influence fertility in opposing ways. We incorporated female employment as a control variable in our baseline regression model (see Section 2). We see no direct reasons as to why age dependency would affect individual choice on child bearing. Finally, the reciprocal of wage efficiency contains on the main part labor efficiency and on the smaller part the distribution of GDP between capital and labor. Both labor efficiency and the share between capital and labor in and of themselves may be just distantly linked to fertility decisions.

So instead of the pension rate, we use the pension index (P^x) as one of the explanatory variables. We define P^x as the product of coverage and the replacement rate, or

$$(7) \quad P^x = (B/O)(b/w).$$

⁴ For the decomposition of pension costs compared to the GDP see Simonovits (2003:8).

Coverage and the replacement rate are indices in which maturity of the pension system most clearly comes into sight. Coverage increases as more and more individuals reaching retirement age fulfill the necessary minimum contributory period entitling them to pensions. While replacement, in addition to other factors, increases if the contributory period increases compared to the productive period of the life-cycle and this also results in higher entry pensions.

Table 3. Aggregate time-series estimates of the effect of family benefits index (FB^x) and the pension index (P^x) on the total fertility rate (TFR) next year. All variables enter the regression in differences, TFR , FB^x and P^x , crude marriage rate and infant mortality in log differences.

Dependent variable: $\Delta \ln TFR_t$	Equation (1 ^x)	Equation (4 ^x)	Equation (5 ^x)
$\Delta \ln FB^x_{t-1}$	0.220 (0.052)	0.186 (0.056)	0.233 (0.047)
$\Delta \ln P^x_{t-1}$	-0.124 (0.101)	-0.135 (0.102)	-0.165 (0.092)
$\Delta \ln TFR_{t-1}$	<u>0.246</u> (0.120)	<u>0.249</u> (0.120)	-0.037 (0.125)
Constant	Yes	Yes ^a	Yes ^a
Current $\Delta \ln FB^x_t$ and $\Delta \ln P^x_t$	No	Yes ^a	No
Other controls	No	No	Yes
Number of observations	52	52	52
R ²	0.42	0.45	0.60
Serial Correlation LM test (χ^2) (p-value)	0.51 (0.78)	1.30 (0.52)	2.73 (0.25)
Long-run effect of FB^x	0.292 (0.098)	0.353 (0.112)	0.225 (0.060)
Long-run effect of P^x	-0.164 (0.129)	-0.268 (0.175)	-0.159 (0.085)

Note: Regression coefficients are *significant at the 1% level*, *significant at the 5% level* and *significant at the 10% level*.

^a Variable(s) entered but not significant at the 5% level.

By analogy, the other explanatory variable, the family benefit index (FB^x), is also composed as a product of coverage and the replacement rate (total public spending on child and family related cash transfers per children under 18, as a share of net real wages).

We tested the models specified in Section 2 with P^x and FB^x as explanatory variables. Estimated effects of family benefits are similar to the baseline, GDP-normalized version. This indicates that the estimated effects are identified by the numerator, not the denominator, providing yet another strong support for the causal interpretation for family benefits affecting fertility.

The same does not hold, however, for pensions. If the pension index replaced the pension rate in the regressions, the estimated effects become a lot smaller. Although in our baseline specification (equation 5 with the control variables) the pension index is still statistically significant at the 10 percent level, the relationship is clearly weaker. This calls for further research in order to understand the interrelationship between the pension system, its various elements, and GDP on fertility.

Conclusion

Building on a theoretical background of inter-linked intergenerational transfers, we tested the fertility effects of family benefits and pensions. We restricted our analysis to public cash transfers but a revised analysis may include in-kind transfers, such as education, health care and long-term care, which have characteristic age-profiles and consequently can also be considered intergenerational transfers.

We found a strong and robust effect of family benefits on fertility. We also observed a strong but less robust effect of pensions on fertility. In addition, we found that much of the estimated pension effect may prove to be a result of measurement error. In particular, employing a standard decomposition technique we showed that the usually applied explanatory variable, the share of pension expenditures in the GDP, cleaned from components that are irrelevant for the individual fertility choice, have much weaker relation with the dependent variable, the total fertility rate. This calls for further research on the effects under examination in a richer data set of micro-data or panel data.

Our finding is not necessary an ultimate blow for the “child as investment good” theory. It is partly because the effect of pensions on fertility did not disappear altogether and partly that the investment nature of child raising is not fully taken into account in the model.

Alternative uses of children, such as, for instance, child labor and private law enforcement in a traditional society with feeble public security, are missing from our model. A proper measurement of such factors may improve the fitness of the model and clean the individual effects of intergenerational transfers on fertility.

4. References

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