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FACULTY OF MATHEMATICS, PHYSICS AND INFORMATICS

Consumption and Income in Slovakia

Dissertation

Matúš Senaj

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To my family

Table of Contests

Abstract

The first chapter focuses on human capital and housing in Slovakia during the economic reforms of the last two decades. We compare households who entered the labor market before and after the economic reforms in 1989. On the one hand, we study the returns to education by different labor market cohorts using household consumption surveys. On the other hand, we analyze the determinants of housing wealth and its impact on consumption. We show that old cohorts are characterized by lower returns to human capital and consumption levels, but higher housing wealth.

The second chapter studies impacts of disposable income and financial wealth on aggregate household consumption. Results confirm that not only disposable income but also financial wealth has significant impact on consumption. We show that the most appropriate proxy for wealth is the sum of monetary aggregate M2 and assets invested in mutual funds. We also investigate the effects of interest rates, consumer confidence index and further relevant variables. It turns outthat these variables are not significant in the consumption function. The second main objective of this work is to evaluate three different consumption forecasting approaches. We show that the most accurate in sample and out of sample forecasts originate from a vector error correction model with exogenous variables.

Finally, we focus on the concept of wage flexibility which is especially important for economic policies after the Slovak euro adoption. The aim of this study is to assess the extent of wage rigidities in Slovakia. We first reproduce Holden and Wulfsberg (2007) approach with data on industrial level drawn from recent decade and we include both old and new EU Member States countries. In case of Slovakia, however, it is difficult to interpret results obtained from sectoral data. Therefore, we turn to micro-approach and apply slightly modified methodology on the company level data. The estimated extent of both nominal and real rigidity is relatively small. Conclusion that hourly compensations are rather flexible supports the decision of euro adoption in 2009.

Abstrakt

Prvá kapitola je zameraná na ľudský kapitál a nehnuteľný majetok na Slovensku v období ekonomických zmien za posledné dve dekády. Porovnávame domácnosti, ktoré vstúpili na trh práce pred a po zmenách v roku 1989. Na jednej strane študujeme výnosy zo vzdelania pre rôzne kohorty domácností. Na druhej strane analyzujeme determinanty nehnuteľného bohatstva a jeho vplyv na spotrebu domácností. Ukážeme, že staršie domácnosti majú nižšie výnosy z ľudského kapitálu, no zároveň majú vyšší nehnuteľný majetok.

Druhá kapitola sa zaoberá odhadom spotrebnej funkcie domácností v SR. Skúmali sme vplyv disponibilného príjmu a bohatstva na konečnú spotrebu domácností v období rokov 1996 - 2005. Zistili sme, že okrem disponibilného príjmu aj finančné bohatstvo významne vplýva na spotrebu a najlepším proxy pre bohatstvo je súčet peňažného agregátu M2 a aktív v podielových fondoch. Tiež sme skúmali vplyv úrokových sadzieb, indexu spotrebiteľskej dôvery a ďalších premenných, tie však nemali významný vplyv na spotrebu. Venovali sme sa aj testovaniu troch rôznych prístupov pri prognózovaní spotreby. Zistili sme, že najpresnejšie in sample a out of sample prognózy sme získali použitím vector error correction modelu s exogénnymi premennými.

Cieľom tretej kapitoly je odhadnúť mieru mzdových nepružností na Slovensku. Koncept mzdovej pružnosti je dôležitý pre menovú politiku, obzvlášť z hľadiska dopadov prijatia eura na Slovensku. Najprv reprodukujeme metódu Holdena a Wulfsberga (2007) s údajmi na úrovni priemyslu čerpanými z uplynulého desaťročia, pričom pracujeme s údajmi starých aj nových členských štátov EÚ. V prípade Slovenska je však ťažké interpretovať výsledky získané zo sektorových údajov, keďže vzorka obsahuje príliš málo negatívnych pozorovaní. Preto sme použili prístup založený na mikro údajoch a aplikovali na údajoch na úrovni podnikov mierne upravenú metodológiu. Odhadovaná miera nominálnych nepružností je pomerne nízka. Záver, že hodinové kompenzácie sú na Slovensku relatívne flexibilné, podporuje rozhodnutie prijať euro v roku 2009.

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The first chapter of the thesis is based on my joint research with Jarko Fidrmuc. We benefited from comments and suggestions made by Jiri Sláčalek and Markus Eller. We also appreciate comments by Katarína Kotovová, Marek Radvanský, Menbere Workie Tiruneh, Ken Iwatsubo and the participants of the international conference on Regional Disparities in Smolenice in 2010. We would also like to thank František Cár, Vladimír Čičmanec and Helena Súkeníková who kindly provided us with the data.

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Introduction

This work is a collection of three empirical papers focusing on consumption and income in Slovakia. In all three papers we apply various econometrical and statistical methods to analyze economic data on both micro and macro level. The first and the second chapter study consumption as household consumption accounts for more than half of gross domestic product in most countries. Last chapter assesses the income flexibility in Slovakia.

The first chapter is based on my joint research with Jarko Fidrmuc titled “Human Capital, Consumption, and Housing Wealth in Transition”, which will be published as a working paper of the National Bank of Slovakia. The paper deals with individual data on Slovak households. We focus mainly on human capital and housing wealth in Slovakia during the last two decades. We compare households that entered the labour market before and after the economic reforms in 1989. The economic reforms in Central and Eastern Europe bring benefits mainly to young and highly qualified people. Young cohorts receive an access to education without former political restrictions, open career opportunities in growing economies and a possibility to travel and work abroad. In contrast, old households had to bear the welfare costs of economic reforms. Because restructuring was associated with the destruction of non-efficient jobs, people often experienced shorter or longer periods of unemployment. Furthermore, their skills and work experience acquired in large state enterprises were often less demanded in market economy. However, we argue that the early cohorts are not necessarily worse off after the reforms.

The second chapter aims at consumption function from the macro perspective. Using the aggregate data, we estimate the elasticities on income and financial wealth in the consumption function. We also investigate the effects of interest rates, consumer confidence index and other relevant variables on consumption. Furthermore, we try to find the best proxy for financial wealth. Moreover, we evaluate three different consumption forecasting approaches in terms of forecast accuracy.

Finally, the focus in third chapter is on the concept of wage flexibility which is especially important for economic policies after the Slovak euro adoption. This chapter is based on my joint research with Pavel Gertler titled „Downward Wage Rigidities in Slovakia", which was published in Czech Economic Review. Wage flexibility is an important concept for monetary policy. It enters into central banks´ thinking about optimum currency areas as well as into its thinking about optimum level of inflation and consequent setting of inflation target. Knowing the extent of wage flexibility is therefore important in any monetary environment; while having own monetary policy or being a part of larger monetary union. The aim of this chapter is to assess the extent of wage rigidities in Slovakia. We analyze industrial data as well as company level data.

1 Human Capital, Consumption, and Housing Wealth[1]

Human capital and physical capital accumulation belong to the most important determinants of growth (Barro, 1991, Levine and Renelt, 1992). The importance of the latter factor was stressed also by central planning countries in Eastern Europe. By contrast, central planning had an ambiguous relationship to the importance of human capital. On the one hand, Eastern European countries focused on basic and technical education (Fischer et al., 1997). On the other hand, human and social sciences were underdeveloped, in some cases even prohibited and persecuted. Similarly, top quality research was often concentrated on high priority secret military project with low spillovers to other sectors.

The early literature on restructuring often stressed the misallocation of resources in central planning economies. In general, industry and especially heavy and military industry received too much weight while services were underdeveloped. As a result, the first years of transition were characterized by shrinkages of industrial production and a fast expansion of the service sector. Less attention was paid to the structure of human capital and the implications of the past education policies on households in Eastern Europe. We try to fill this gap in the literature. In particular, we compare human capital equipment and especially the returns to education for cohorts which finished their education before and after 1989.

Moreover, we look also at other characteristics of household wealth. While people faced restrictions in their accumulation of human capital before the changes of the 1990s, the socialist system with intensive social subsidies provided some important benefits to the population as well. Although the majority of them were short-run benefits, they included also easy and cheap access to social housing. By contrast, residential construction went down during the first years of reforms. Supply declines caused excess demand more recently, which resulted in a housing bubble in several countries (Hlaváček and Komárek, 2009).

[1] This chapter is based on my joint research with Jarko Fidrmuc, which will be published as a working paper of the National Bank of Slovakia.

Different past conditions in the education sector and housing market caused persistent differences between households in Eastern Europe. In general, old households are characterized by restricted access to higher education but by higher endowment of housing stocks. In contrast, young generation enjoys a better access to competitive education. However, the supply of housing to these households is constrained by imperfect housing markets. We use this time variation in access to education and housing in order to estimate their effects on private consumption, which stands as a proxy for welfare.

Overall, we can see a similar tradeoff that we observed for firms in the first stage of economic reforms. High investment of old households is associated with insufficient and low-quality human capital equipment, while young households have higher human capital equipment but low access to physical capital (housing). These features may have two long-term implications for Eastern European countries. First, it is more difficult to change capital structure of households than to support restructuring of the firms. People often do not re-qualify until they are forced to do so by extreme events, including for example long-term unemployment. Often people even loose their previous qualification and move to less skilled and worse paid occupations.[2] Moreover, it is hard to acquire specific skills in older age. Fidrmuc and Fidrmuc (2009) for example document that there is still a strong division line in language skills between Western and Eastern European countries, the latter region being less able to communicate in all foreign languages (with the exception of Russian).

Second, different factor equipment has important implication for political economy and political stability in Eastern European countries. Economic reforms have introduced welfare gains but also losses. Households have been diversely affected by these changes. This often motivates their political behavior (Fidrmuc, 2000). Denisova, Eller and Zhuravskaya (2010) show that the attitudes towards transition and the role of state are different for Russian respondents with different characteristics. Older and less educated respondents are more likely to view critically the economic reforms and favor a more important role for the state in the economy. Therefore, it is important that no population groups are excluded from growth and welfare improvements.

The economic reforms in Central and Eastern Europe bring benefits mainly to young and highly qualified people. Young cohorts receive an access to education without former political restrictions, open career opportunities in growing economies and a possibility to

[2] Campos and Dabušinkas (2009) show that the majority of occupational changes in Estonia between 1990 and 1994 were towards sectors with lower wage levels.

travel and work abroad. In contrast, old households had to bear the welfare costs of economic reforms. Because restructuring was associated with the destruction of non-efficient jobs, people often experienced shorter or longer periods of unemployment. Furthermore, their skills and work experience acquired in large state enterprises were often less demanded in market economy. However, we argue that the early cohorts are not necessarily worse off after the reforms. We identify the physical capital equipment (housing) as an important source of their wealth. Since the housing sector was heavily subsidized during the former regime, those people own apartments and houses more likely than younger households. Several measures of economic reforms have aimed to improve the economic situation of the incumbent households at the expenses of future developments. This includes the sale of housing at low costs to resident population, but also the voucher privatization which was introduced in several Eastern European countries. Our results indicate that this was actually important in order to equalize the welfare effects of economic reforms on different cohorts.

We concentrate on Slovakia, because this country is an example of a fragile liberal democracy. Slovak economic policy belongs to the forerunners among the new member states,[3] nevertheless, the political developments faced regular populist and nationalist trends and flashbacks (Malová and Miháliková, 2002). Correspondingly, specific redistributive policies (e.g. voucher privatization and sales of public apartments to tenants) were actively targeted especially by liberal parties. We view these policies as a part of compensation policies which target the political economy equilibrium in the country.[4]

Using a detailed dataset on households in Slovakia, we show that the returns to education are significantly different between cohorts with human capital acquired before and after the reforms. We use household income survey data which are more appropriate than wage data used in other papers because they cover all sources of income. This may be important especially in countries with a high share of informal economy. Moreover, we can illustrate significant differences between housing wealth of both cohorts, which were not addressed in the earlier literature. We use information on quality and statute of housing in order to impute the value of real estate owned by the households. We show that housing quality compensates old households at least partially for income losses due to their low returns on

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[3] Slovakia introduced a major reform of taxation including the flat tax in 2004 (Moore, 2005). It joined the European Union in 2004 and the euro area in 2009.

[4] Similar arguments in favor of voucher privatization are presented by Roland and Verdier (1994).

education. Overall, it is difficult to identify winners and losers of economic reforms, because they cannot be attributed to demographic groups. We argue that this contributed to stabilization of the political system and allowed significant reform progress in Slovakia.

The paper is structured as follows. Next part reviews the previous literature on returns to education and housing wealth in Eastern Europe. Section 1.2 describes our data sets on consumption expenditures and housing wealth in Slovakia. Part 3 provides descriptive analysis of households' income, housing and consumption. Empirical results are discussed in section 4. Section 5 concludes and generalizes our results from Slovakia for further countries in Eastern Europe.

1.1 Literature overview

In general, central planning countries tried to reduce all sources of inequality (Orazem and Vodpivec, 1995, Campos and Coricelli, 2002). Correspondingly, there was a tendency to equalize wages for all jobs, possibly excluding several priority areas as e.g. the heavy industry.[5] As a result, returns to education were negligible in all central planning countries. Münich, Svejnar, and Terrell (2005) present a deep comparison of returns to human capital under the communist regime and during the transition to the market economy. They find that the returns to education were extremely low before 1989, but increased already during the first years of transition. These results are largely similar to earlier estimations for the Czech Republic and in Slovakia presented by Chase (1998), for Romania by Andrén, Earle and Sapatoru (2005), or for Slovenia (Orazem and Vodpivec, 1995). More recently, higher returns to education were reported by Newell and Socha (2007) for Poland. These results are confirmed by Fleisher, Sabirianova and Wang (2005), who document in a meta analysis that the average returns to schooling doubled between 1990 and 2002 in transition economies (including China). Orlowski and Riphahn (2009) contributes to the literature by studying the returns to tenure and experience in East and West Germany. They find that the returns to experience are lower in East Germany, which is probably a result of economic transition when the experience of some workers become obsolete.

Several authors address the suitability of human capital achieved at communist education system. In particular, previous authors discuss several ways how education attained before economic reforms may be less valued in a market economy. First, it is often argued that

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5 Münich, Svejnar, and Terrell (2005) present the wage grid applied in the Czech Republic for industry, heavy industry and public sector.

education was concentrated to areas (e.g. rocket science) which are less demanded in market economies (Campos and Dabušinkas, 2009). Second, important soft skills in marketing and management may be missing (Campos and Coricelli, 2002). Third, however, the quality of education could be worse because of external shocks which were not related to economic transition. Card and DiNardo (2002) show that low experience with computer usage causes a negative wage premium in the US. Finally, low education premium for tenured employees may be perpetuated in their later wage profile if wage setting practices are using former income as a negotiation base for later wages (Andrén, Earle and Sapatoru, 2005).

Several authors test whether the returns to education attained before economic reforms are lower than returns to market-type education. Contrary to the initial expectations, the previous analysis show statistically insignificant difference between returns to education attained before and after 1990. Münich, Svejnar, and Terrell (2005) compare the marginal returns of a year of education completed either before or after 1990. Somewhat surprisingly, they found lower returns to education completed during the economic reforms. Andrén, Earle and Sapatoru (2005) also find no statistical difference between education acquired before and after economic reforms.

However, these results may be influenced by a short number of observations with post-communist education (about 14% of sample used by Münich, Svejnar, and Terrell (2005), which represents about 320 employees). It may be influenced also by adverse labor market developments during the first year of transition (reduction of employment and increase of youth unemployment). Moreover, the quality of education could worsen during the reforms in Romania as argued by Andrén, Earle and Sapatoru (2005). Campos and Coricelli (2002) note that human capital indicators (e.g. enrollment rates) declined in all transition economies including also Central Europe. Finally, the previous authors consider the wage premium of an additional year of education, while the impact may be different on employees with basic, secondary or high education.

There are only a few analysis of housing wealth in Eastern Europe. Early studies pointed at the importance of housing as a part of non-wage benefits. In the planned economies, municipalities and firms were often made responsible for providing social services to employees and population (Tsenkova and Turner, 2004, Juurikkala and Lazareva, 2006). State enterprises used the non-wage benefits to attract employees. Thus, the enforced equalization of income was at least partly compensated by non-wage benefits, and housing played a key role in this respect. Berger, Blomquist and Sabirianova Peter (2008) show that the Russian employees were compensated for differences in regional living quality through better access to housing.

Contrary to other areas, economic reforms did not target the distortions in the housing market. Low supply resulted in soaring housing prices. Égert and Mihaljek (2007) show that housing price grew similarly as income. By contrast, Hlaváček and Komárek (2009) find several periods of housing bubbles (in particular, in 2002-2003 and 2007-2008) in the Czech Republic. The differential development of housing markets before and after 1990 implies that households have largely different stocks of the housing wealth.

For developed economies, numerous papers estimate the housing wealth effect on consumption based either on household level data as well as on aggregate data. They usually report that marginal propensity to consume with respect to wealth change is up to 0.10. This means that exogenous increase in the value of the house of 1 percent leads to an increase of household consumption by 0.10 percentage points. Farinha (2009) uses micro-level data collected in Portugal in 2006 and 2007. This survey confirms that housing wealth represents the most important asset in the portfolio of household. She estimates the elasticity of consumption with respect to wealth between 0.04 and 0.05 for households. For Italy, Grant and Peltonen (2008) report marginal propensity to consume above 0.08 percent of housing wealth. They also find that the marginal propensity to consume is higher for old household (0.15) whereas it is statistically insignificant for young households. In Spain, (2006) finds that the marginal propensity to consume is 0.02 percent of housing wealth of the main residence. For aggregate consumption, Carrol, Otszuka and Slacalek (2011) report the wealth effect between 4 and 10 cents per $1 change in housing wealth in U.S. data. Using a simple life-cycle model, Attanasio, Leicester and Wakefield (2011) prove that house price shock should have larger effect on the consumption of older households. Moreover, they conclude that consumption of younger households tends to respond more to income shocks. Contrasting to the previous findings, Calomiris, Longhofer and Miles (2009) claim that housing wealth has a small and insignificant effect on consumption in the US if the control for the endogeneity bias caused by correlation between housing wealth and the permanent income. So far, however, there are no comparable studies for new member states of the EU.

1.2 Data Description

In our repeated cross-sectional analysis we merge two different data sources for households in Slovakia. First, we use data on household income and consumption from the household expenditure survey (HES), which is conducted annually by the Statistical Office of the Slovak Republic. Second, we use data on house prices which is published quarterly by the National Bank of Slovakia. Both datasets are available for the period from 2004 to 2009. Thus our data are not influenced by the flat tax reform in 2004 (Moore, 2005). The datasets may be influenced only marginally by the accession to the European Union, which occurred in May 2004. We do not deflate the nominal data; instead we include time effects in the estimated equations.

The HES is collected since late 1950s. However, due to several important methodological changes implemented in the household survey in 2004, we cannot use the previous household surveys for the analysis.[6] The survey provides data on structure of income and expenditures of Slovak households. The sample size of the HES is approximately 4,700 households every year. Since 2004, the Statistical Office of the Slovak Republic uses random sampling. The selected sample is representative not only on a country level but also at the regional level. The basic sampling unit is a private household[7] created by one or more persons who (i) live together in the same dwelling and (ii) share the costs of living.

Given the data quality, the household expenditure surveys are used commonly for regular analysis of the household expenditures, the computation of general population weights, calculation of consumer price indices, and as a main source for final private consumption entering the National Account System (Byfuglien, 2006). The methodological changes in 2004 improved significantly representativeness of the surveys. Nevertheless, the sample may not include very rich and very poor households (Deaton, 2005, Carraro, 2006). Another restriction is the lack of time dimension, which means that we can not trace a behavior of a specific household over the sampling period.[8]

The household expenditure survey provides detailed information on housing quality, but it does not include the market value of the housing facilities. In order to impute the housing wealth, we use our second dataset which comes from the survey of residential property prices[9] conducted jointly by the National Bank of Slovakia and the National Association of

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[6] Before the 2004 vawe, the HES used to have smaller sample size (only about 1600 households) and significant part of Slovak households were not surveyed. Particularly those households with unemployed person, disabled person or single mother as a head of the household.

[7] Collective households (for example hospitals, cloisters, and prisons) are excluded from the survey.

[8] An alternative household survey, EU SILC, includes aslo a panel component. However, the EU SILC survey does not cover information on household consumption.

[9] Detailed information on the survey can be found in Cár (2006).

Real Estate Offices. We consider the housing wealth equals to the market value of the house because we do not have the information on mortgage[10]. Based on the region, location and the number of rooms of the property, we distinguish between 120 different types of a flat or a house. From the household expenditure survey, we know the type of property the household occupy. Moreover, from the survey of residential property prices we draw the prices of every group of dwelling out of 120 mentioned earlier[11]. In particular, we use square meter prices of the property with given characteristics from this database. We impute the value of housing wealth by multiplying the size of the flat or house with the corresponding square meter price. This imputation adds the information on housing wealth to the original HES database.

1.3 Descriptive statistics

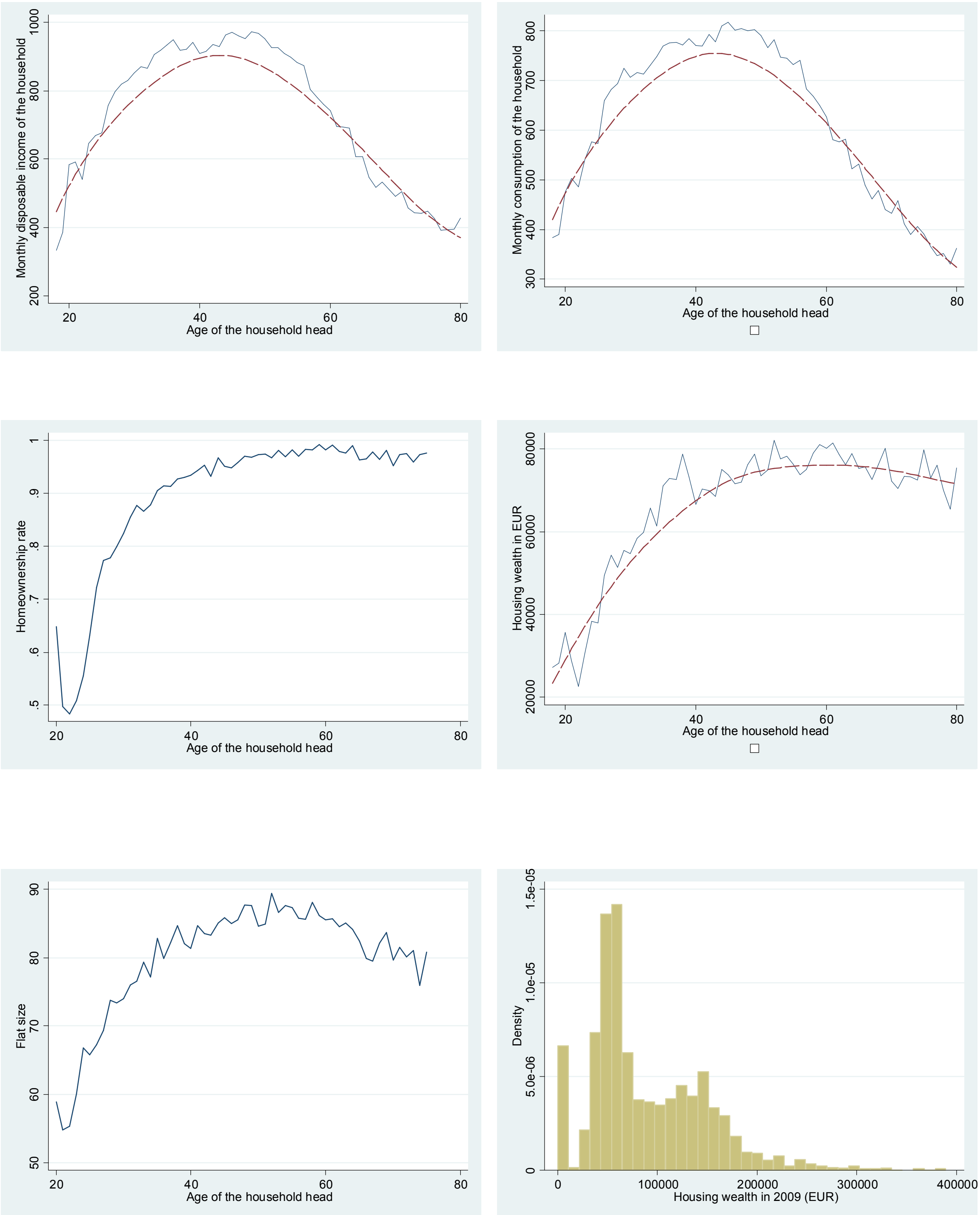
A first insight into the database offers several interesting features of household's disposable income. As can be seen in the figure 1.1, age-income profile has a nonlinear shape. It increases to the age of 50 of the household head (HH). The income of older families sharply decreases. Moreover, the variability reaches its highest levels between the ages of 50 and 60. The income profiles correspond with those of consumption. However, the age-consumption profile decreases faster than income. This means that older households are more intended to save.

Since the housing wealth is the most important component of private capital we focus our attention on housing. Contrary to the previous findings, older households own more housing wealth. In general, former communist countries are known as countries with very high home-ownership rate due to the mass privatization of former state rental housing (Lux 2004, Edgar, Filipovic and Dandolova, 2007). This is also the case of Slovakia. Figure 1.1 depicts also the pattern of ownership rate in our sample and displays the snapshot of dependency between the age and home ownership. The ownership rate starts at 50 percent for the youngest households. Then it grows to the values higher than 95 percent. More than 95 percent of households older than 50 years own their residence.

[10] The share of households with mortgage is about 10 per cent in Slovakia.

[11] For example, prices range from 418 euro to 5130 euro per square meter in 2009.

Figure 1.1 Descriptive statistics by age cohorts



Source: authors' calculation.

Note: Consumption, income and housing wealth are deflated (in 2005 prices).

Very similar pattern can be found in age - housing wealth profile. Housing wealth rises over the life cycle. However, it is slightly declining for the oldest households. This is a natural feature. We can expect that older households usually occupy smaller flats. What can be surprising is an average house price per square meter for different age groups. Although, properties of older households are smaller, they are more valuable mainly because of better location of their apartments and houses. The last graph shows the histogram of housing wealth in Slovakia. In our sample, we have more than 7 per cent of households that do not own any real property - they are included in the first bin.

1.4 Empirical analysis

In this section we proceed in three steps. Firstly, we start our analysis with a focus on human capital of the households and its impact on the disposable income. In the second step, we turn our attention to housing wealth and look for the determinants of the residential property prices. Finally, we use both indicators (disposable income and housing wealth) as determinants of the consumption function.

1.4.1 Income determinants

We start with the estimation of the income equation. The depended variable is defined as monthly household income (in logs), although the traditional Mincer model of earnings (Mincer, 1974) analyzes individual income instead of a household of several persons. This reflects that the next sections analyze housing wealth and consumption of households.

Table 1.1 presents estimations of household income determinants. The first specification (labeled Income 1) includes standard demographic indicators as gender, age, education, which are defined for the principal earner (household head). Since, the age-income profile has nonlinear shape; we include two age variables in the regression. Variable age is defined as age of the household head divided by 10 and age2 stands for $age^2/100$. Moreover, we control if the household head is widow, widower, divorced or single (we call this group as single parent). Furthermore, the number of household members and number of children are included in order to capture the size of the households. All variables are highly significant and they have the expected signs. The households with highest education have higher income by 18.5 percent as compared to secondary education (omitted category). Households with female principle earner have income lower by about 4.8 percent, but single households with children have income lower even by 23.1 percent. We also include the level of partner's education. Similarly to negative gender effects, we find that the partner's education yields significantly lower returns to human capital. For example, the households with university educated household head have higher income by

18.5 percent higher, however, the household with the same level of education of his/her partner earn more by 13.7 percent. The negative differential of the secondary earner is visible also for primary education, where it accounts to 0.9 percent.

The next specification (Income 2) extends the analysis by a cohort differential in age-income profile of the Slovak households. In particular, this specification includes also Cohort90, which is a dummy variable that equals to one if the principal earner entered the labor market before 1990 (in other words, if his or her age was more than 25 years in 1989).[12] These households received their complete education before economic reforms. Such specification shows that income of households belonging to the Cohort 90 is lower by 7.1 percent compared to the younger households.

However, cohort differentials may correspond to several factors including previous earning profile (see discussion in section 2). Therefore, we compare the returns to human capital for different levels of education in the last specification. As compared to secondary education achieved after 1990, which was selected as a base category, employees with university degree receive lower incomes if they entered the labor market before 1990, while income for respondents with basic education is actually higher for those, who entered the labor market before 1990. This contradicts alternative explanations of income differentials (e.g. wage persistence).

These results are highly robust to various sensitivity tests. Following Chase (1998) we include years of education and years of experience in the disposable income equation. Given expected length of education for various levels, we define years of education as

$$\textit{years of education} = (\textit{basic} * 8 + \textit{secondary} * 4 + \textit{university} * 5) / 10 .$$

Additionally, we define the potential experience indicator, which represents an additional characteristic of human capital. The variable experience is defined, as it is common in the literature:

$$\textit{potential experience\_1} = (\textit{age} - \textit{years of education} - 6) / 10 .$$

$$\textit{potential experience\_2} = (\textit{age} - \textit{years of education} - 6)^2 / 100$$

[12] This threshold was confirmed in a version of Chow structural break test for the household's age between 20 and 30 in 1990. The age of 25 years in 1990 was associated with the highest *t*-statistics.

Table 1.1 OLS estimates of disposable income equation

	Income 1	*Income 2*	*Income 3*	*Income 4*	*Income 5*	*Income 6*
Age	0.095***	0.161***	0.163***			
Age2	-0.016***	-0.020***	-0.021***			
Number of adult members	0.271***	0.276***	0.275***	0.244***	0.292***	0.275***
Number of children	0.058***	0.052***	0.053***	0.030***	0.064***	0.049***
Female	-0.048***	-0.048***	-0.049***	-0.076***	-0.019	-0.039***
Single parent	-0.230***	-0.226***	-0.227***	0.090	0.119***	-0.227***
Education primary	-0.122***	-0.122***	-0.328***			
Education tertiary	0.185***	0.185***	0.187***			
Partner's education primary	-0.131***	-0.131***	-0.128***			
Partner's education tertiary	0.137***	0.136***	0.135***			
Cohort 90		-0.071***				
Cohort 90 x primary education			0.148***			
Cohort 90 x secondary education			-0.074***			
Cohort 90 x tertiary education			-0.076***			
Years of education				0.488***	0.265***	
Years of partner's education				0.261***	0.304***	
Potential experience				0.377***	-0.242**	
Potential experience squared				-0.103***	-0.050	
Years of education before 1990						0.415***
Years of education after 1990						0.494***
Potential experience before 1990_1						-0.173***
Potential experience after 1990_1						0.096***
Potential experience before 1990_2						0.041***
Potential experience after 1990_2						0.036***
Constant	9.115***	8.958***	8.961***	8.135***	8.197***	8.604***
Sample	All households	All households	All households	Cohort of younger h.	Cohort of older h.	All households
No. of obs.	27 650	27 650	27 650	8 674	19 288	27 962
R^2	0.641	0.642	0.643	0.464	0.665	0.637

*Note: */**/*** depicts statistical significance of the coefficients at 10/5/1 % level of significance respectively. Regional and time effects are not reported. Source: authors' calculation.*

We split the sample for both cohorts according to the age of the principal earner. Two columns (Income 4 and Income 5) of Table 1.1 presents regression of household incomes on years of education, years of potential experience, experience squared and other variables capturing households' characteristics. Similarly to the previous results, the estimated coefficient on years of education is significantly lower among the older households, what confirms our previous findings. Moreover, the effect of experience is also lower in the older cohort.

Finally, we compute years of education and years of potential experience achieved before and after 1990 for every household. The last column of Table 1.1 summarizes the estimated coefficients. Not surprisingly, it turns out that the education attained after 1990 are significantly better rewarded compared to the education from transformed schooling system after 1990. Thus, our sensitivity analysis confirms that human capital obtained before economic reforms yields lower returns than human capital attained more recently.

1.4.2 Determinants of housing wealth

Under central planning, housing was under tight state control. Majority of dwellings [13] were owned directly by the state or by state enterprises. Lux (2004) classifies the following four types of housing: state rental flats, rental flats owned by state enterprises, cooperative rental flats and privately owned family houses. The mass privatization of former state rental housing began in the early 1990s. Due to privatization, a lot of families became owners of flats they had occupied. Moreover, they had to pay far less than the market price for such flats. Thus, older families acquired their housing wealth relatively cheaply and they are often equipped with a higher physical capital.

The estimation of determinants of housing wealth has to reflect that the sample includes households with zero reported housing wealth. In this case, the OLS approach will not yield consistent estimates mainly because the censored sample is not representative of the population (Cameron and Trivedi, 2009). Therefore, we apply two alternative approaches. The Tobit model proposed by Tobin (1958) reflects the truncation of the data sample. However, it does not consider that households with housing wealth are likely to be different from the remaining households. The Heckman selection model reflects also the selection bias in the housing wealth equation.

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[13] Due to data limitation, we assume only dwellings the families live in. So we do not consider another properties owned by the family.

The Tobit model is used when the depended variable is not always observed but the vector of covariates is fully observed. This is also our case because housing wealth is observed only if the household owns the dwelling. Therefore, we try to apply the Tobit model on housing wealth. The model is specified for unobserved (latent) variable model:

$$y^* = x\ \beta + u$$

Where, housing wealth is proxied as

$$housing\ wealth = \begin{cases} y^* & if\ \ y^* > L \\ L & otherwise \end{cases}.$$

The variable x stands for vector of covariates and L is a censoring point. The model is left censored.

Explanatory variables include disposable income, number of family members, dummy variable for single parent households and also dummy for households belonging to Cohort 90. The estimated coefficients for all variables are in line with the standard expectations. Moreover, the coefficient on Cohort 90 is also positive and highly significant. The impact is actually large also in economic terms. Older households, who received housing under the central planning, own property valued nearly 50% more than younger households.

As many authors point out, the Tobit model is very sensitive to proper specification. The estimates are inconsistent if the errors are not normally distributed or if they are not homoskedastic. Therefore, we use a selection model, which might be more appropriate approach for the estimation of the housing wealth. The Heckman selection model offers a more general modeling approach than the Tobit model. It represents a bivariate sample selection model. In general, the model includes a selection equation for home ownership, (ho):

$$ho = \begin{cases} 1\ \ if\ \ z\gamma + u_2 > 0 \\ 0\ \ if\ \ z\gamma + u_2 \le 0 \end{cases}.$$

The outcome or regression equation for housing wealth has linear form. The dependent variable housing wealth (hw) is observed only if ho equals to one. Thus

$$hw = \begin{cases} x\beta + u_1 & if\ \ z\gamma + u_2 > 0 \\ - & otherwise \end{cases},$$

where x stands for vector of covariates for outcome equation, z represents vector of covariates for selection equations. The Heckman selection model can be estimated by two different methods. First, a two-step estimation approach estimates the outcome and selection equation in two steps. Second, both equations can be estimated simultaneously by a maximum likelihood, which is generally considered to be more efficient and robust. The Table 1.2 presents results of both methods which supports the robustness of our results.

Table 1.2 Tobit and Heckman selection model estimates of housing wealth

	Housing 1	Housing 2	Housing 3	Housing 4	Housing 5	Housing 6
	Tobit model	Tobit model	Heckman model – ML	Heckman model -2STEP	Heckman model - ML	Heckman model -2STEP
Disposable income	0.159***	0.145***	0.061***	0.043***	0.061***	0.054***
Number of family members	0.031***	0.078***	0.050***	0.056***	0.063***	0.062***
Single-parent	-0.149***	-0.137***	-0.058***	-0.045***	-0.060***	-0.058***
Cohort 90		0.448***			0.133***	0.093***
Constant			13.434***	13.619***	13.286***	13.395***
			Homeownership – selection model			
Disposable income			0.370***	0.336***	0.360***	0.336***
Single parent			-0.199***	-0.238***	-0.206***	-0.238***
Number of family members			0.042	0.019	0.034	0.019
Age			1.023***	1.093***	1.058***	1.093***
Age2			-0.067***	-0.076***	-0.071***	-0.076***
Cohort 90			0.041	-0.000	-0.031	-0.000
Constant			-5.301***	-4.931***	-5.178***	-4.931***
No. of observations	26 300	26 300	27 965	27 965	27 965	27 965
No. of censored obs.	1 674	1 674	1674	1674	1674	1674
rho			-0.447	-1.000	-0.275	-0.701

*Note: */**/*** depicts statistical significance of the coefficients at 10/5/1 % level of significance respectively. Regional and time effects are not reported.*

Source: authors' calculation.

We include the same set of explanatory variables as in the previous Tobit estimation. Moreover, the selection equation includes age and age squared of the principal earner, which are used as the exclusion restriction14. This reflects the assumption that households are more likely to become home owners with their age. Once obtained, households stay in the same housing for a relatively long period of time. Therefore, the value of housing does not change again with age (it is actually insignificant if included in outcome equation). Nearly all variables keep expected signs in the selection equation and age variables are

[14] Because of the identification condition, we use age in the selection equation only. Therefore, it does not appear in the regression equation.

highly significant. By contrast, number of members of the family as well as the dummy variable for Cohort 90 is insignificant and they do not have any effect on home ownership.
The outcome equation contains the disposable income, which as expected has a positive sign. Families with higher income live in properties that are more valuable. The coefficient on number of family members has also positive sign indicating that larger families own dwellings that are larger and thus more valuable. All specifications confirm that households belonging to the Cohort 90 own more valuable housing wealth compared to younger households. Year dummies capture the development of house prices compared to the omitted year 2004. We include also regional dummies to control for different regional prices of properties, which keep expected results. For example, households in the capital city, Bratislava, are less likely to own their housing, but if so, it is more expensive than in other regions.

1.4.3 Consumption function

So far, we have shown on the one hand that families which entered the labor market before the transition have significantly lower disposable income. On the other hand, those families own higher housing wealth. However, neither income nor housing wealth individually describe welfare sufficiently. Therefore, we merge both channels in an estimation of household consumption, which may be taken as a better proxy of households' welfare [15]. In particular, we study the impact of income and housing wealth on consumption. By and large, the household consumption forms more than 50 per cent of output. Such importance is another reason to study the impact of wealth on consumption.
Using the Durbin-Wu-Hausman test, we strongly reject the hypothesis that disposable income is exogenous variable in the consumption function. Since the OLS estimation would be biased in this case, we prefer an instrumental variable approach to estimate the consumption function. In order to present robust estimates, we perform two-stage OLS together with GMM estimation. The crucial point here is the selection of proper set of instruments. We test the exogeneity of instruments by Hansen J-statistics. Given its results, we choose two variables as proper instruments, namely a dummy for households with female principal earner and households living in a town. These instruments describe the background of the household. We use gender income differences (lower income for female

[15] In order to keep our analysis simple, we do not discuss possible implications of future income streams and time preferences between the households.

principal earner) which are well documented in the literature (see e.g. Weichselbaumer and Winter-Ebmer, 2005 and 2007) and confirmed by previous estimations. Moreover, we suppose that income in towns is higher, which reflects better earning opportunities in urban centers.[16] We do not include cohort dummy variable into instruments due to possible endogeneity. The reported Hansen J-statistics confirms that our instruments are valid (see table 1.3). The same group of instruments is used in two-stage OLS and also in GMM estimation. In order to present robust results we have utilized two stage OLS and Generalized methods of moments approach. Using both methods, we have estimated two types of consumption function. They both include disposable income, housing wealth and dummy for primary and tertiary education. The only difference is that the second and forth specifications are supplemented with the variable Cohort 90. Therefore, we present four columns in the table 1.3. In all cases, the estimated coefficients on disposable income are more or less the same. The marginal propensity to consume is around 0.87. It turns out that housing wealth effect on consumption is insignificant. Actually, this result may reflect a low degree of development of credit markets in Slovakia. Thus it is generally impossible to use mortgage equity withdrawals.

Table 1.3 IV estimates of consumption function

	Consumption 1	Consumption 2	Consumption 3	Consumption 4
	2SLS	2SLS	GMM	GMM
Disposable income	0.874***	0.862***	0.872***	0.860***
Housing wealth	-0.001	0.001	-0.001	0.001
Primary education	-0.113***	-0.106***	-0.114***	-0.107***
Tertiary education	0.015*	0.017**	0.014*	0.016**
Constant	1.149***	1.280***	1.170***	1.296***
Cohort 90		-0.047***		-0.046***
Number of observations	27 965	27 965	27 965	27 965
Centered R^2	0.620	0.623	0.620	0.624
Hansen J statistic	0.164	0.157	0.164	0.157
Hansen p-value	0.686	0.692	0.686	0.692

*Note: */**/*** depicts statistical significance of the coefficients at 10/5/1 % level of significance respectively. Time effects are not reported. Disposable income is instrumented by dummies for female principal earner and for location in urban centers.*

Source: authors' calculation.

[16] Both instrumental variables are correctly signed and highly significant in the fist state equation. F-Statistics of excluded instruments is with values of 820 and 885 well above the recomended threshold of 10.

We can see that early labor market cohorts have a lower consumption level than younger cohorts by almost 5 percent. Nevertheless, consumption of older households would be higher if rents for better housing were properly imputed. In the following example, we try to evaluate the size and the impact of imputed rents. In Table 1.2, we estimate that value of housing of older households is higher by 9 to 13 percent. Given the annual rent to price ratio of 5.5 percent (Global Property Guide, 2012) and the average consumption and house price this implies that imputed rent for older households is increases their consumption by 5 to 7 percent of consumption. Therefore, the differences in consumption levels between the cohorts are rather negligible if rent imputation is considered. Thus we can conclude that the negative income level is largely compensated at the level of consumption.

Table 1.4 shows the estimates of consumption function for different levels of education and for different age groups. Both characteristics apply to principal earner. The main difference between the three groups of estimates is the coefficient on disposable income. Low educated families have the highest marginal propensity to consume. As before, the housing wealth effect is insignificant across all three groups.

Table 1.4 Consumption function by education and age categories

	Primary education	Secondary education	Tertiary education	Age<=30	30<Age<=40	40<=Age<50	Age>50
Disposable income	0.938***	0.849***	0.847***	0.865***	0.834***	0.839***	0.889***
Housing wealth	0.000	0.001	-0.000	0.002	0.001	-0.001	0.003**
Cohort 90	-0.090**	-0.049***	-0.029**				
Constant	0.550**	1.403***	1.463***	1.223***	1.548***	1.513***	0.935***
No. of observations	3324	20839	3802	1907	5179	6572	14307
Centered R^2	0.579	0.580	0.512	0.426	0.457	0.504	0.644
Hansen J statistic	0.895	0.060	0.220	0.055	0.021	1.105	3.518
Hansen p-val	0.344	0.806	0.639	0.814	0.885	0.293	0.061

*Note: */**/*** depicts statistical significance of the coefficients at 10/5/1 % level of significance respectively. Time effects are not reported. GMM estimator is used.*

Source: authors' calculation.

Finally, we present the estimated coefficients for four age categories. First group includes younger families, where the principal earners are younger than 30 years old. Principal earners between 30 and 40 and between 40 and 50 comprise to the second and third column, respectively. The last column represents families older than 50 years. In general, the estimates are more or less comparable across all age groups. The marginal propensity to consume is a bit lower in the second and third age group. However, it is higher among the youngest and the oldest households. The coefficient on housing wealth is insignificant in first three groups. Although it is significant among the oldest families, the value of the coefficient is almost zero.

1.5 Conclusions

Central planning in Eastern Europe put a large weight on material production and investment. Behind specialization of these countries on heavy industry, we show that these preferences were strongly anchored also in provision of human capital and physical goods to population. The access to education was strongly controlled and education was focused on areas enjoying preferences by central planning. By contrast, the countries made a significant effort to satisfy basic needs of population, including also affordable housing for everybody.

Also nearly 25 years after the beginning of economic reforms and transition in Eastern Europe, the former preferences are clearly visible in the endowment of population by human capital and physical capital. Using household expenditure surveys in Slovakia, we demonstrate significant differences between cohorts entering the labor market before and after 1990. On the one hand, returns to human capital are lower for education acquired before market reforms. On the other hand, the early labor market cohorts enjoyed also easier access to housing. We find that both effects seem to counteract each other to a significant degree. Older employees face lower returns to human capital, which lowers their disposable income. Keeping other effects unchanged, this would result in negative implications on their welfare. However, older households enjoyed also a preferential access to housing. Although the effects are difficult to quantify exactly, the magnitude of both effects, and their expected variability among individuals, lead to conclusion that it is difficult to identify winners and losers of transition, at least in the example of Slovakia. The importance of both effects describes how fragile is the political support for economic reforms in Eastern Europe. We argue that mass privatization programs (voucher privatization, but even more privatization of housing to incumbent tenants) played an important role for ensuring political support during economic reforms.

Behind political economy considerations, we provide several findings with regard to income, wealth, and consumption determinants at the household level. Household income reflects education level of its members. However, we find also significant gender and regional differences. Disposable income is an important determinant of housing ownership as well as of its value. Finally, household consumption is determined by available income. Instrumental estimations of income effects imply marginal propensity to consume which is below 1 and it varies strongly for household categories. Finally, housing wealth is not entering directly the consumption function, which reflects low availability of financial instruments in Eastern European countries.

Appendix

Table 1.5 Summary statistics

	No. of observations	Mean	Std. Dev.	Min	Max
all households					
Household consumption	27 377	672	378	44	3272
Disposable income	27 377	796	438	13	4035
Housing wealth	27 377	71484	61738	0	1325641
younger households					
Household consumption	8 678	743	357	78	3110
Disposable income	8 678	894	407	101	3731
Housing wealth	8 678	65379	63399	0	1128155
cohort 90 households					
Household consumption	18 699	640	383	44	3272
Disposable income	18 699	750	445	13	4035
Housing wealth	18 699	74318	60745	0	1325641

Table 1.6 first stage estimates of consumption function

	Consumption 1	Consumption 2	Consumption 3	Consumption 4
Dependent variable: disposable income				
Housing wealth	0.01***	0.02***	0.01***	0.02***
Primary education	-0.36***	-0.34***	-0.36***	-0.34***
Tertiary education	0.20***	0.20***	0.20***	0.20***
Cohort 90		-0.09***		-0.09***
Female household head	0.05***	0.05***	0.14***	0.05***
Type of town	0.14***	0.14***	0.22***	0.14***
Constant	0.22***	0.21***	0.32***	0.21***
Number of observations	27 962	27 962	27 962	27 962
Prob > F	0.00	0.00	0.00	0.00
Partial R2	0.17	0.17	0.17	0.17

*Note: */**/*** depicts statistical significance of the coefficients at 10/5/1 % level of significance respectively. Time effects are not reported. Disposable income is instrumented by dummies for female principal earner and for location in urban centers.*

Source: authors' calculation.

2 Consumption Function Estimate and Consumption Forecasting[17]

In most countries, final household consumption accounts for more than half of GDP. The consumption function is thus one of the most frequently discussed topics in theoretical as well as empirical literature. From the point of view of the central bank, it is important to know the elasticities of disposable income or other variables and to be able to forecast the future development of the consumption, as the personal consumption expenditure affects the demand-pull inflation.

The aim of the paper is twofold. First, to estimate the coefficients of consumption function in Slovakia. Second, to find a proper way of forecasting of the path of household consumption.

In the first part of the article, we are going to present the best-known works in the field of modern consumption theory. In 1935, Keynes introduced his absolute income theory, in which he says that consumption was a function of disposable income. Later empirical research, however, came to the conclusion that this relationship does not provide a sufficient explanation for the behaviour of aggregate data.

The life cycle theory assumes that consumption makes up a constant part of the present value of lifetime income. Foundations of the theory were laid by Modigliani and Brumberg (1954). This theory says that the propensity to consume is lower with young households than with households of older persons who spend their savings. Aggregate demand thus depends not only on income and wealth, but also on demographic changes.

An extension of the Keynes theory has been presented by Friedman (1957). His permanent income hypothesis divides consumption and income into a permanent and a transitory element and includes future income expectations. If an individual considers a change of his

[17] This chapter is based on my research paper titled "Consumption Function Estimate and Consumption Forecast: the Case of Slovak Republic", which was published in Ekonomicky casopis (Journal of Economics), 56(1), pages 3-21.

income transitory, he has no reason to change his consumption habits. On the other hand, if he finds out that the income change is of permanent nature, he simultaneously adjusts his consumption. According to Friedman's theory, the key determinant of consumption is the real wealth of the consumer, not his current real disposable income.

The credibility of the permanent income theory with rational expectations is weakened by the presence of obstacles for borrowing – the "liquidity constraints". The existence of such constraints causes consumption to be determined more by present than future income and the interest rate not to play a major role in consumption decisions (e.g. Hall and Mishkin (1982) or Zeldes (1989)).

The said theoretical works are taken up by empirical research, which has dealt with estimating the consumption function. Takala (1995) says that the inclusion of a variable representing wealth entails more stable consumption function estimates. Côté and Johnson (1998) have added the consumer attitudes index to the short-term variables. This step has increased the proportion of the explained variance of the dependent variable and has enhanced the accuracy of the consumption prediction. Bredin a Cuthbertson (2001) have modeled a consumption function for the Czech Republic for the years 1993 to 1995. The authors have found a long-run equilibrium between consumption, real wages and inflation. They have also found out that the size of the income effect depended significantly on the decision to include the variable unemployment rate in the estimated model. Selected aspects of consumption in the Czech Republic have been also treated by Artl et al. (2002). In their work, they describe the co-integration between seasonally adjusted time series of real consumption and of real disposable income. The vector error correction model for household consumption in Spain was designed by Marínez-Carrascal and del Rio (2004), who have analyzed the impact of loans granted to households on the consumption of the households. The effect of financial and housing wealth has been analyzed e.g. by Bover (2006) and Sierminska and Takhtamanova (2007).

The paper has the following structure. In the next part, we summarize the methodology. Third section describes the data. In the section four, we discuss the obtained results. Finally, last part concludes.

2.1 Methodology

This section briefly describes the methodology. Firstly, we focus on the Johansen approach and its application on non stationary time series. Secondly, we present the criteria for evaluation of forecasts.

2.1.1 The estimate of consumption function

The selection of the methodology used to estimate the consumption function has been influenced by the presence of non-stationarity in our data sample. Therefore, the Johansen approach based on the vector autoregression methodology (VAR) is used to estimate the consumption function. This methodology is used, similarly as the Engel-Granger approach, when working with non-stationary time series. It is also suitable for predictions at a short-term horizon (Pindyck and Rubinfeld, 1991). Gujarati (2003) says that in many cases forecasts obtained using this method is better than those obtained from complex simultaneous models. We will first briefly introduce the VAR models and then show how to convert the VAR model to a VEC model.

As an illustration, we will introduce a simple VAR model, consisting of the two endogenous variables y_1 and y_2 (both are non-stationary, type I(1)). For the sake of simplicity, we will assume two lags:

$$\begin{aligned} y_{1,t} &= \mu_1 + a_{1,1}y_{1,t-1} + a_{1,2}y_{2,t-1} + b_{1,1}y_{1,t-2} + b_{1,2}y_{2,t-2} + \varepsilon_{1,t} \\ y_{2,t} &= \mu_2 + a_{2,1}y_{1,t-1} + a_{2,2}y_{2,t-1} + b_{2,1}y_{1,t-2} + b_{2,2}y_{2,t-2} + \varepsilon_{2,t}\ , \end{aligned} \qquad (1)$$

which corresponds to the following in vector notation:

$$y_t = \mu + A_1 y_{t-1} + A_2 y_{t-2} + \varepsilon_t\ . \qquad (2)$$

The equation (2) contains a vector of endogenous variables (y_t), a vector of constants (μ), matrices of the coefficients of lagged variables (A1 and A2) and a vector of deviations (ε_t).

The VAR model notation can be transcribed to an error correction form. The resulting model is called vector error correction (VEC):

$$\Delta y_t = \mu + \Pi y_{t-1} + \Gamma_1 \Delta y_{t-1} + \varepsilon_t\ . \qquad (3)$$

Based on the rank of the estimated matrix Π, we can decide whether the variables y_1 and y_2 are co-integrated as well as decide on the number of co-integration vectors. The number of co-integration vectors is always lower than the number of endogenous variables. In this case, the size of the matrix is 2x2. If the rank of the matrix is two, then y_1 and y_2 are stationary and the cointegration is not an issue. If the matrix Π has only zero values, then Δy_t depends only on Δy_{t-1} and vectors y_1 and y_2 are not cointegrated. Finally, if the rank of the matrix is one, then both vectors y_1 and y_2 cointegrated. Basically, this is how the test of number of cointegrated vectors works. In this paper, we use so called trace test which tests the rank of the matrix Π.

The Johansen approach employs the maximum likelihood methodology for estimation error correction forms of VAR models.

2.1.2 Methodology of forecast evaluation

One of the main goals of the paper is to find the most suitable way of consumption forecasting. In order to measure the predictive ability of the suggested models, we construct the dynamic ex post forecast, which is then confronted with the actual values of consumption. As evaluation criteria, we use the following measures.

SSF *(Sum of squares of forecast error)* is defined as a sum of the squares of the deviations of the actual data from the forecasted values:

$$SSF = \sum_{i=1}^{p} \left(Y_{T+i} - Y_{T+i}^{f} \right)^2 , \tag{4}$$

where Y_{T+i}^{f} is predicted value and Y_{T+i} refers to the actual value. The number of forecasted periods is denoted as *p*.

An index RMS *(Root-mean-square error)* refers to the previous one:

$$RMS = \sqrt{\frac{1}{p} \sum_{i=1}^{p} \left(Y_{T+i} - Y_{T+i}^{f} \right)^2} . \tag{5}$$

We put the main emphasis on an index PMSP (*Root mean square percent error),* because it enable us to measure the correctness of forecasts of different time series. The value of this index is normalized with the time series under inspection.

$$RMSP = \sqrt{\frac{1}{p} \sum_{i=1}^{p} \left(\frac{Y_{T+i} - Y_{T+i}^{f}}{Y_{T+i}} \right)^2} * 100 \quad \text{[in per cent]}. \tag{6}$$

2.2 Data

We use quarterly data for the time period 1996 - 2005 to estimate the consumption function. In order to compare out-of-sample consumption forecasts with the actual values , we do not include year 2006 into the estimates,. All time series are converted to real values using the consumer price index and they have been seasonally adjusted. We realize that in terms of the general conception it would be better to use the consumption deflator, but we are not convinced about the reliability of the deflator data. In addition, the said deflator can be changed during data revision.

The modelled variable in this material is abbreviated as *c* and represents the final consumption of households. The main data source is the system of national accounts which is set up by the Statistical Office of the Slovak Republic (in m. SKK).

The analyzed time period can be characterized as a period of positive economic development in Slovakia. The real household consumption and also the disposable income grow over the whole period with the exception in the years 2000 and 2003 (see Figure 2.1) when the consumer prices grows faster. As a result the real consumption and income slightly decrease.

Figure 2.1 Household consumption (in logs)

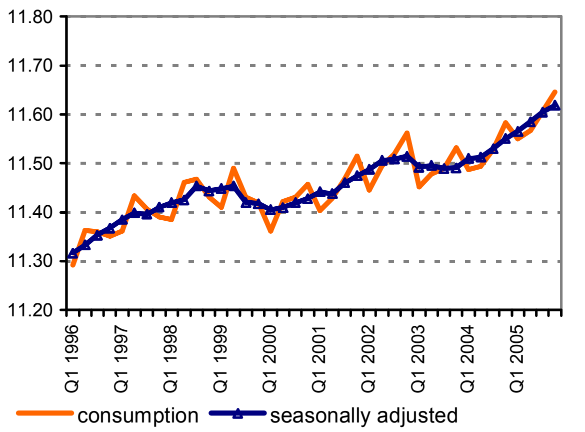
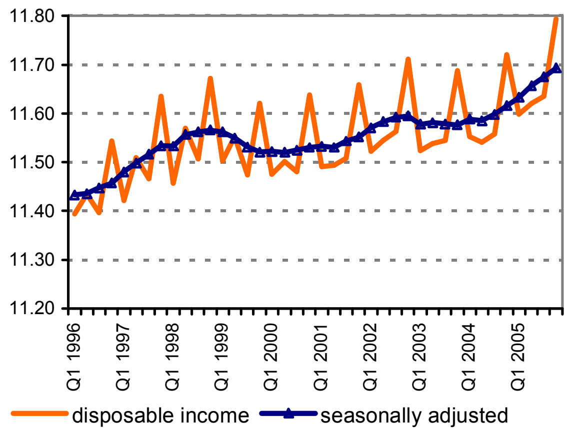


Figure 2.2 Disposable income (in logs)



Source: Statistical Office of the Slovak republic.
Note: In constant prices (1995).

Following Côté and Johnson (1998) and Singh (2004), we use variables reflecting both income elasticity and the wealth effect.

Household income is most frequently represented by gross disposable income (*y*). This time series comes from the national accounts and it is expressed in millions of SKK. We expect the disposable income to have a positive impact on consumption with elasticity between 0 and 1.

The selection of the variable for reflecting the financial wealth effect is somewhat more complicated. Financial wealth is usually approximated by the monetary aggregate (M2) in the literature (Artl et al., 2001, Filáček, 1999). Some authors subtract bank loans from the M2 aggregate (eg. Berdin and Cuthberson, 2001). Another one uses quasi money as a proxy (Singh, 2004). In the case of Slovakia, probably the best solution would be to use data regarding the financial assets of households, which are also included in the national accounts. A drawback of this approach is the fact that only annual data is available, which renders the use of the data impossible in this analysis. Therefore, we offer and test four variables as financial wealth proxies. The first two variables are the monetary aggregate M0, i.e. currency in circulation, and the monetary aggregate M2, including cash savings. We later extend the monetary aggregate M2 to include household assets in mutual funds (PF) , which have become an increasingly important part of population savings. As the last proxy, we test the sum of quasi money, i.e. the difference between the aggregates M2 and M1, and assets in mutual funds. The data source of monetary aggregates is the National bank of Slovakia. We expect the wealth elasticity of consumption is positive and smaller than one.

According to published empirical works, we identify several variables that can affect consumption. One of them is the level of interest rates. We have used the BRIBOR interbank rates with a maturity of 1, 3 and 6 months, as well as rates on loans to households. Higher interest rates make saving more attractive and are likely to reduce also the demand for loans and this can subsequently be reflected in a drop of final consumption. Viewed from this standpoint, the interest rates have a negative impact on household consumption. On the other hand, higher interest rates cause an increase in the financial wealth of households, and if the wealth effect is positive, the impact of interest rates on consumption is positive, as well.

Figure 2.3 Proxy for wealth (monetary aggregate M0)

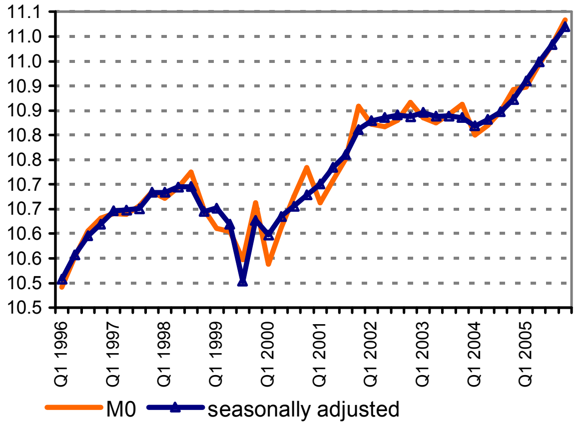


Figure 2.4 Proxy for wealth (monetary aggregate M2)

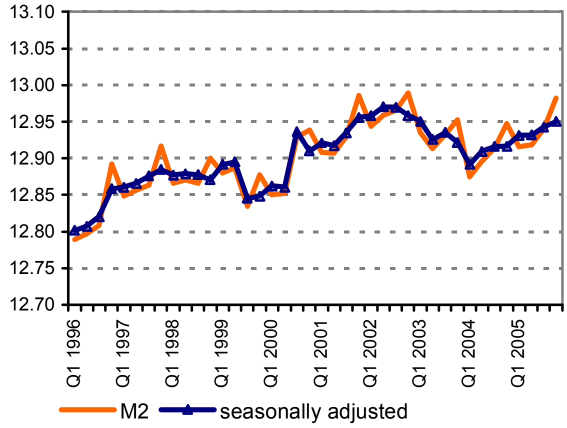


Figure 2.5 Proxy for wealth (monetary aggregate M2+mutual funds)

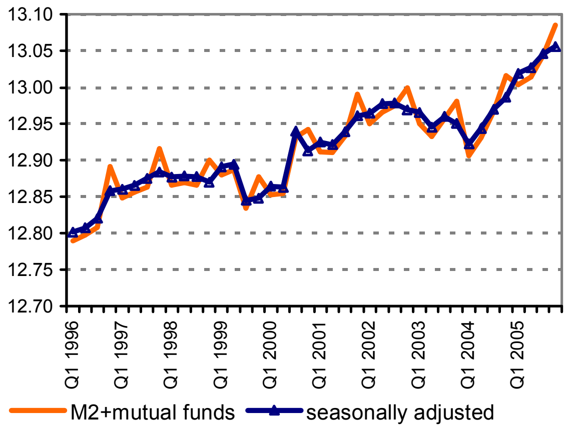
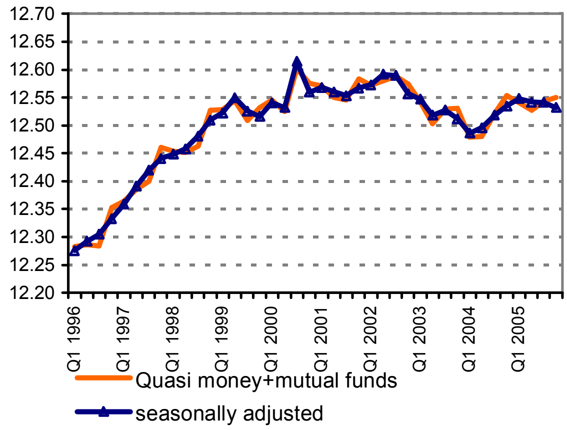


Figure 2.6 Proxy for wealth (quasi money+mutual funds)



Source: National bank of Slovakia.

Note: In constant prices (1995).

The growth of the employment (denoted "zam") has a positive impact on consumption, because the higher the number of employees, the income of households is also higher. However, this effect is uncertain, as the impact of the growth of the employment rate is already included in the higher disposable income of the households. On the other hand, an employed person has better expectations regarding future income than an unemployed person, which, according to the life cycle theory, increases current consumption.

Following Côté and Johnson (1998), we try to include the consumer confidence indicator (ISD) in the estimated equations. We expect that positive outlooks of the consumers might indicate the increase in consumption.

Table 2.1 Summary statistics

	C	Y	M0	M2	M2+MF	QM+MF	ZAM	MZDA	ISD
	[mil. Sk]	[mil. Sk]	[mil. Sk]	[mil. Sk]	[mil. Sk]	[mld. Sk]	[tis.osôb]	[Sk]	[body]
Mean	95 028	104 321	46 470	401 086	408 898	268 837	2 166	8 342	72
Maximum	111 203	119 739	61 101	429 558	468 091	301 021	2 241	9 081	90
Minimum	82 138	92 326	36 449	362 795	362 795	214 356	2 091	7 670	61
St. dev.	6 743	6 103	6 029	17 597	26 272	21 690	44	346	8
No. of obs.	40	40	40	40	40	40	40	40	24

Source: NBS, SO SR, author's calculation.

Note: In 1995 prices.

The data is deflated by the index of consumer prices and it is seasonally adjusted with the Tramo/Seats approach. The statistical properties of the variables are presented in Table 2.1.

2.3 Results

In the next part we summarize the results obtained from VEC analysis of Slovak macro variables.

2.3.1 Consumption function estimate

We expect the variables we use to be non-stationary, and also the formal tests conducted confirm the validity of this hypothesis. For stationarity testing, we use an Augmented Dickey-Fuller test (ADF) and a Kwitakowski-Philips-Schmidt-Shin test (KPSS). The difference between them is that the null hypothesis of the ADF test assumes the presence of a unit root while the null hypothesis of the KPSS test assumes stationarity of the examined time series. Identical results of both tests therefore considerably decrease the possibility of rejection of a true hypothesis. The Table 2.2 shows the test statistics and critical values of both tests applied on the time series of consumption, disposable income, proxy for financial wealth and real wage.

The results of both tests are unambiguous for all variables except quasi money and tell us that the time series are not stationary. Their differences, however, are stationary. Therefore we consider them integrated of order one, which we denote as I(1).

Table 2.2 Stationarity tests

	test statistics		critical value (5%)	
	ADF	KPSS	ADF	KPSS
				levels
c	0.44	0.75	-2.95	0,46
y	0.14	0.70	-2.95	0,46
a (M0)	-0.39	0.72	-2.94	0,46
a (M2)	-2.20	0.61	-2.94	0,46
a (M2+PF)	0.70	0.76	-2.94	0,46
a (QM+PF)	-3.15	0.49	-2.94	0,46
wage	0,36	0.71	-2.92	0.46
	1-st. differencies			
c	-3.32	0.12	-2.95	0,46
y	-3.15	0.13	-2.94	0,46
a (M0)	-7.05	0.10	-2.94	0,46
a (M2)	-7.40	0.15	-2.93	0,46
a (M2+PF)	-6.85	0.05	-2.94	0,46
a (QM+PF)	-6.47	0.58	-2.94	0,46
wage	-6,23	0.19	-2.92	0.46

Source: author's calculation.

Since we model the relationships between non-stationary variables, their mutual cointegration has to be tested. Therefore, we apply the Johansen co-integration test to all four presented long-term equations. The test confirms the existence of a stationary combination of these variables. Thus, we can say that consumption, disposable income and wealth are cointegrated. In addition, only one cointegration vector in all cases. We use three lags, what means that current consumption is affected by the values from the four preceding quarters.

Using the vector error correction model, we model the relationship between three variables. Long-term coefficients obtained from this estimate represent an estimate of the parameters of the consumption function, in particular elasticity on income (α) and the elasticity on financial wealth (β).

$$v_t = \left(c_{t-1} + const + \alpha y_{t-1} + \beta a_{t-1}\right)\rho + \gamma_0 + \sum_{i=1}^{p} \gamma_i^T v_{t-i} + u_t \, , \tag{7}$$

where v_t is the vector of the first differences of endogenous variables, ρ is the vector of adjustment coefficients, γ_i are vectors of estimated coefficients under lags of endogenous variables, γ_0, *const* stands for vector of constants and u_t is the vector of deviations. The number of lags is limited by the value p in this expression. As mentioned above, p equals to three in our estimates. For the sake of completeness, let us express the said relationships also by means of a formal notation:

$$v_t = \begin{pmatrix} \Delta c_t \\ \Delta y_t \\ \Delta a_t \end{pmatrix}, \ \rho = \begin{pmatrix} \rho_1 \\ \rho_2 \\ \rho_3 \end{pmatrix}, \ u_t = \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ u_{3,t} \end{pmatrix}.$$

Estimates of the long run equilibrium between consumption, disposable income and wealth are presented in Table 2.3. We consider four alternatives with various definitions of wealth. In the first model, wealth is approximated by currency in circulation M0 and by the monetary aggregate M2 in the second model. In the third model, we extend M2 with assets invested in mutual funds. In the last model, the sum of quasi money and assets invested in mutual funds is used.

Table 2.3 Estimates of consumption function

Dependent variable: c_{t-1}

		model 1	model 2	model 3	model 4
const		-1.19	-5.15	-3.75	-2.03
α		0.95	1.20	0.93	1.17
		(-12.97)	(-17.26)	(-13.23)	(-17.31)
β	$m0_{t-1}$	0.16			
		(-4.85)			
β	$m2_{t-1}$		0.21		
			(-2.78)		
β	$m2_{t-1} + pf_{t-1}$			0.34	
				(-5.46)	
β	$qm_{t-1} + pf_{t-1}$				$1.1 \cdot 10^{-6}$
					$(2.6 \cdot 10^{-5})$
Number of lags		3	3	3	3
Number of cointegrating vectors[18]		1	1	1	1

Source: Author's calculation.

Note: t-statistics are in parenthesis.

[18] According to Johansen procedure. Results of the test are presented in tables 2.18 – 2.21 in the appendix.

Based on the size of the estimated coefficients, we exclude the second model, because income elasticity cannot be higher than 1. For the same reason, we exclude model 4. In addition, if we use the sum of quasi money (QM) and assets in mutual funds as a proxy of wealth, its estimated is no significant. The reason is that the population holds a part of its savings in cash. For example, one third of the respondents say that they save in the form of cash in the Czech Republic (Artl et al., 2001).

According to the statistical properties and economic interpretation, we prefer to select a model featuring the M2 monetary aggregate and mutual funds. This long-term equilibrium quantifies the elasticity on income at 0.93 and the elasticity on financial wealth is 0.34. The signs of both coefficients are in line with our expectations. We assume that the relatively high values of the coefficient in the case of disposable income are primarily due to two factors: the high average propensity to consume (defined as the consumption to disposable income ratio), as well as the low volume of loans granted to Slovak households. This means that the main determinant of household consumption is their disposable income. The financial wealth elasticity is lower.

The long run coefficients are comparable to those estimated for other developed economies. For example, Côté and Johnson (1998) focuses on Canada. They found the income elasticity is 0.98 and financial wealth elasticity is 0.36. According to Lettau and Ludvigson (2001) the coefficients on income and wealth are 0.59 and 0.31, respectively, in the United States. Takala (1995) finds the income elasticity is 0.71 and wealth elasticity is 0.23.

While estimating consumption function, we put primary emphasis on the long run relation between consumption, disposable income and wealth. Short-term relations are shown in Table 2.4. The fact that we work with shorter time series can be seen in higher standard deviations and lower t-statistics of individual coefficients[19]. This causes the problem with their interpretation. Co-integration vector, described above, contains the sum of the monetary aggregate M2 and the assets invested in mutual funds. Adjustment coefficient (ρ) reaches the value of 0.31 and shows the speed of adjustment of consumption to its equilibrium value. That means that if the consumption is unbalanced from the long-term

[19] For comparison, we create a restricted version of VEC model by excluding the first and the second lag of wealth. Then the coefficient of cointegration vector is statistically significant and the estimated coefficients are similar to those presented.

balanced relation perspective, then in one year 78 per cent of this variance should be eliminated.

Table 2.4 The coefficients of dynamic equation

Dependent variable

	coefficient	t-statistics
$Coint_{t-1}$	-0.31	-1.46
γ_0	0.002	1.24
Δc_{t-1}	-0.24	-1.01
Δc_{t-2}	0.54	2.27
Δc_{t-3}	0.58	2.80
Δy_{t-1}	0.70	3.00
Δy_{t-2}	-0.27	-1.08
Δy_{t-3}	-0.63	-2.24
Δa_{t-1}	0.07	0.77
Δa_{t-2}	0.02	0.22
Δa_{t-3}	-0.15	-1.75
Dummy Q1 2003	-0.02	-2.62
Dummy Q1 2004	0.02	2.01

Source: Author's calculation.

Note: Dummy varible Q1 2003 eliminates the effect of price deregulation and dummy Q1 2004 removes the impact of tax structure changes.

2.3.2 Consumption forecasting

In this part we test three possible ways of predicting the consumption. The first one is to create a VEC model, in which all three variables are endogenous. The second possibility is to create three equations for consumption, disposable income and wealth and their interrelation within the model. Finally, we estimate an ARMA model of consumption.

Vector error correction approach

Our VEC models contain three variables – the household consumption, disposable income and proxy variable for wealth. They were calculated in two variants, which vary from each other only in the last variable. In the first case, the wealth is represented by the monetary aggregate M0 – the currency; in the second case the wealth is represented by the aggregate M2 extended with the assets invested in mutual funds.

We start with models estimated without the exogenous variables (tables 2.14 and 2.15 in the attachment). Comparing the adjusted coefficients of determination, the model with the currency seems to be more exact as to the prediction of the household consumption. However, both of the models give rather inexact forecasts. Therefore, we add exogenous variable employment, and two dummy variables (which gain value of one in the first quarter of 2003 and 2004), to the short-term specification. This extension improves the models significantly[20]

Due to the shortness of the time series of the consumer confidence indicator (it has been recorded only since 1999), we do not include this variable as exogenous. However, estimates for a shorter time interval have confirmed that, unlike for Canada (Côté and Johnson, 1998), it does not have a significant impact on the explanation of the short-term deviation of consumption from its equilibrium value in the case of Slovakia. Similarly, an addition of interest rates does not improve the properties of the estimated models. The insignificance of the interest rate coefficient is probably caused by the fact that the interest rate effect on consumption is included in the variable representing wealth both in the long-term and in the short-term relationship. Not even the application of real interest rates has brought about significance of their coefficients.

Table 2.5 Forecasts' evaluation

		Model 1	Model 3	Model 1a	Model 3a
Proxy for wealth		M0	M2+MF	M0	M2+MF
Exogenous variables		Not included	Not included	Included	Included
	c	0.459	0.372	0.698	0.730
Adjusted R^2	*y*	0.449	0.380	0.604	0.547
	a	0.238	0.014	0.345	0.086
	c	2.66 %	2.62 %	1.39 %	0.58 %
RMSP	*y*	2.57 %	2.52 %	0.85 %	0.47 %
	a	3.80 %	3.13 %	5.94 %	1.68 %

Source: Author's calculation.

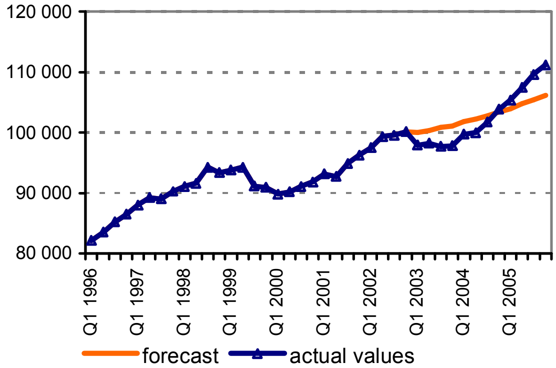
In the table 2.5, we present indicators for evaluation of exactness of dynamic ex post forecasts within the years 2003 - 2005. Here we can observe that the model 1 and also the model 3 have lower values of index RMSP and higher values of the coefficient of determination. This means that they are more exact at the ex post prediction. If we follow only the variable c, then on the basis of the mentioned indicators we prefer to choose

[20] Cointegrating tests are presented in tables 3.22 and 3.23 in the appendix.

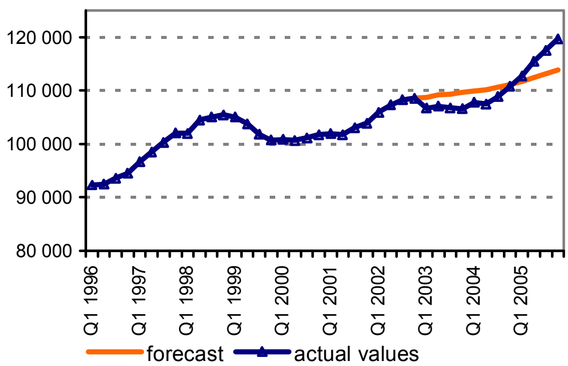
model 3a. Also a graphic comparison indicates that within the years 2003 – 2005, model 3a is the most exact one.

Figure 2.7 Dynamic forecast of the model 1 (time period 2003:1 2005:4)

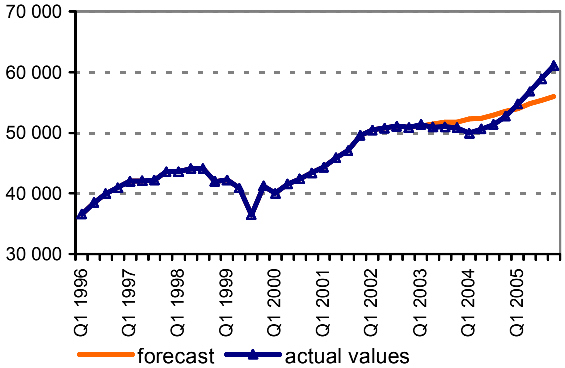
Household consumption



Disposable income



Proxy for wealth (M0)



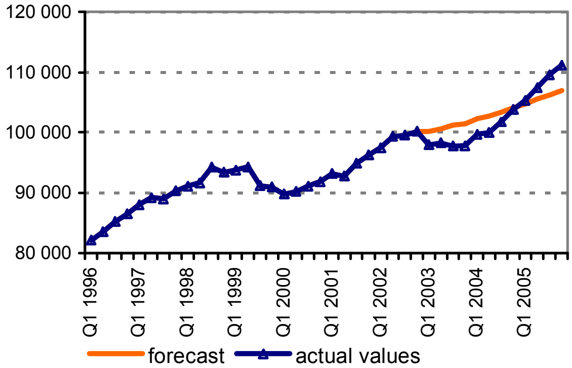
Measures of predictive ability

	c	y	*a (M0)*
SSF	92662592	1.02E+08	56961410
RMS	2778.828	2916.633	2178.712
RMSP	2.66 %	2.57 %	3.80 %

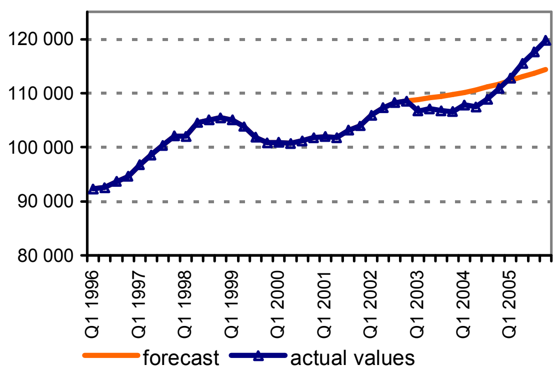
Note: Time period from 1Q 2003 to 4Q 2005.

Figure 2.8 Dynamic forecast of the model 3 (time period 2003:1 2005:4)

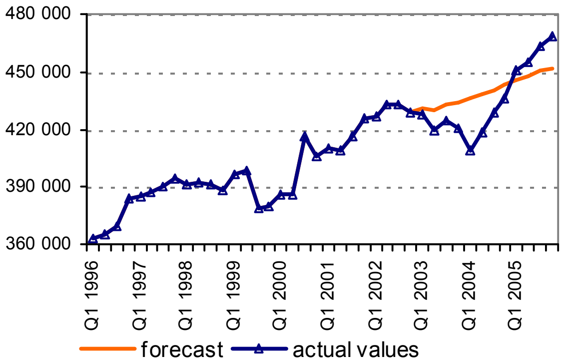
Household consumption



Disposable income



Proxy for wealth (M2+MF)



Measures of predictive ability

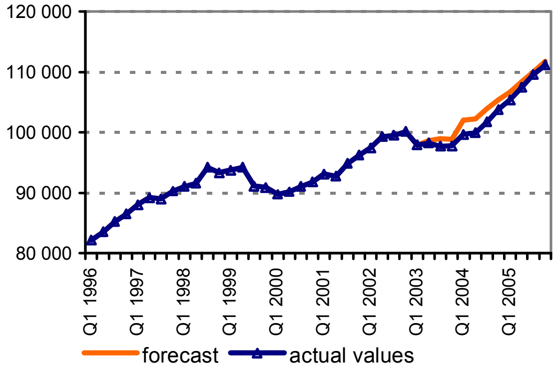
	c	y	*a (M2+pf)*
SSF	86738600	97095019	2.13E+09
RMS	2688.5	2844.5	13333.1
RMSP	2.62 %	2.52 %	3.13 %

Note: Time period from 1Q 2003 to 4Q 2005.

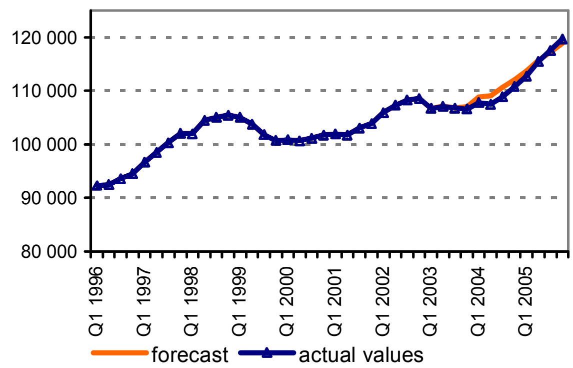
Source: Author's calculation.

Figure 2.9 Dynamic forecast of the model 1a (time period 2003:1 2005:4)

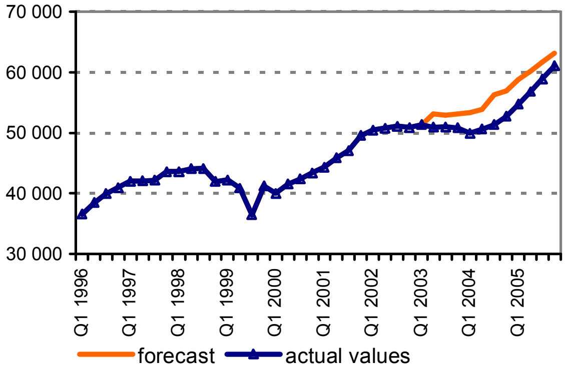
Household consumption



Disposable income



Proxy for wealth (M0)



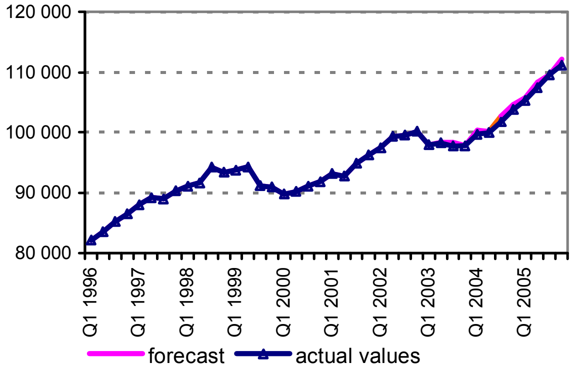
Measures of predictive ability

	c	y	*a (M0)*
SSF	23889671	10360108	$1.19*10^{8}$
RMS	1410.9	929.1	3144.8
RMSP	1.39 %	0.85 %	5.94 %

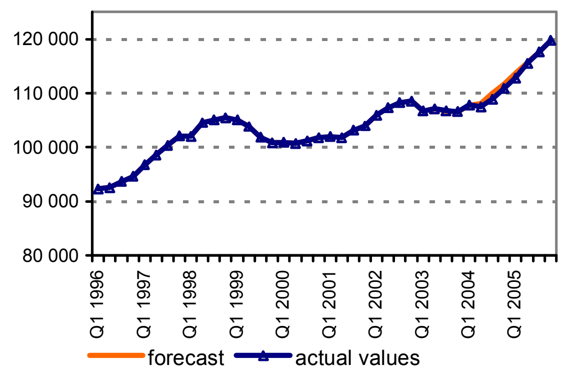
Note: Time period from 1Q 2003 to 4Q 2005.

Figure 2.10 Dynamic forecast of the model 3a (time period 2003:1 2005:4)

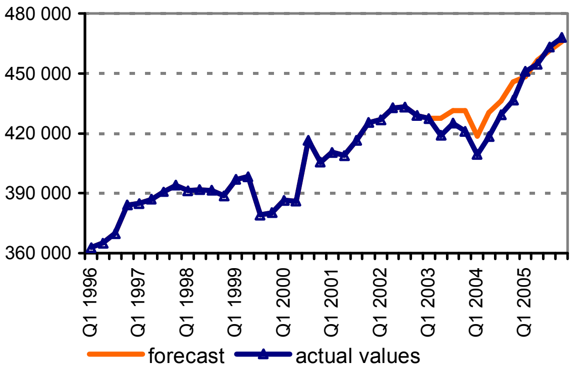
Household consumption



Disposable income



Proxy for wealth (M2+MF)



Measures of predictive ability

	c	y	*a (M2+pf)*
SSF	4448370	3269623	$6.06*10^{8}$
RMS	608.8	521.9	7103.8
RMSP	0.58 %	0.47 %	1.68 %

Note: Time period from 1Q 2003 to 4Q 2005.

Source: Author's calculation.

In order to test the exactness of the model 3a, we compare forecasts the period 2003 – 2005 with the actual data. That means that we compare in sample forecasts. Since, in this study we forecast the household consumption deflated by consumer prices, we cannot directly compare it with the consumption in real prices which is modified by the deflator of consumption. For that reason we adjust the predicted consumption in current prices and we compare it in table 2.6 with the data of the statistical office. It turns out that this model offers more exact in sample predictions as the difference between predicted and real growth in ten cases out of twelve is less than 1 p.b. Moreover, in presented aberrance no systematic trend is present and the prediction ranges narrowly over or under the real value.

Table 2.6 Nominal household consumption growth (in per cent)

	Q1 2003	Q2 2003	Q3 2003	Q4 2003	Q1 2004	Q2 2004	Q3 2004	Q4 2004	Q1 2005	Q2 2005	Q3 2005	Q4 2005
Actual growth	8.4	6.2	6.1	6.3	12.2	9.9	11.8	11.9	9.4	10.2	10.2	10.1
Forecast (in sample)	8.4	6.3	6.8	6.2	12.9	9.9	12.1	12.9	9.1	10.9	9.1	10.1
Difference (forecast/actual value)	0.0	1.6	11.5	-1.6	5.7	0.0	2.5	8.4	-3.2	6.9	-10.8	0.0

Source: author's calculation.

Note: Seasonally unadjusted data is presented.

The estimates of three separate equations

The second way – estimation of the three single equations[21] and then their joining into one model - is often used in Slovakia. A disputative part is to find an equation for wealth, in our case for the sum financial wealth.[22] In general, the estimation of the development of aggregate M2 is difficult in Slovakia. The reason is the inaccessibleness of the date which would represent alternative costs, so-called opportunity costs (Švantner, 2005). The issue of estimation of currency aggregate needs a more detailed analysis and could thus be subjected to further research.

First we will introduce the equation for consumption. That is based on an assumption in part 2.3.1, but we changed the variables explaining short-term deviation from a state of equilibrium. For that reason the coefficients in long-term part are slightly different. To be

[21] The way of constructing these models does not allow the usage of johansen procedure. The cointegration was confirmed by a significant t-statistics on the cointegrating coefficient.

[22] A simple solution is to predict the wealth by ARMA model. However, in this case we obtain very biased forecast of the future development of wealth.

concrete, the elasticity of income has been estimated to 0.81 and the elasticity of wealth to 0.50 ($c_{t-1} = 0.81 * y_{t-1} + 0.50 * a_{t-1} + 4.33$). Short-term dynamic is created by habits (persistence) in consumption; those represent late values of consumption and the change in real disposable income of the population. The following table contains the specification of a dynamic equation.

Table 2.7 Estimates of dynamic equation of consumption

	coefficient	t-statistics
$Coint_{t-1}$	-0.29	-2.50
constant	0.004	2.25
Δc_{t-1}	-0.08	-0.65
Δc_{t-2}	0.14	1.16
Δc_{t-3}	0.16	1.47
Δc_{t-4}	-0.43	-3.93
Δy_t	0.72	5.12
Adjusted R^2 : 0.67		

Source: Author's calculation.

In the second equation, we have used the relation between the real disposable income and real salary and we included these in a co-integrating vector. The estimated long-term salary elasticity reached the value of 0.66, which means that raising real salary of 1 per cent will affect the raise of disposable income of 0.66 per cent. Into the co-integrating vector, we have also added a constant and a linear trend and therefore we may record it as follows: $y_{t-1} = 0.667 * salary_{t-1} + 0.003 * trend + 5.466$. Almost all short-term coefficients are significantly different from zero (table 2.8).

Table 2.8 Estimates of dynamic equation of disp. income

	coefficient	t-statistics
$Coint_{t-1}$	-0.33	-2.95
constant	0.003	1.95
Δy_{t-1}	0.52	3.33
Δy_{t-2}	0.69	3.84
Δy_{t-3}	0.22	1.65
Δy_{t-4}	-0.43	-3.38
$\Delta mzda_t$	0.34	4.70
$\Delta mzda_{t-1}$	-0.10	-0.93
$\Delta mzda_{t-2}$	-0.19	-1.95
Adjusted R^2 : 0.72		

Source: Author's calculation.

As the third, we present an equation used for the forecasting of the financial wealth. The long run relation includes a scale variable GDP in real prices with the estimated coefficient of 0.69.

Table 2.9 Estimates of dynamic equation of wealth

	coefficient	t-statistics
$Coint_{t-1}$	-0.23	-3.05
constant	-0.003	-0.92
Δa_{t-1}	-0.09	-0.81
Δa_{t-3}	0.10	1.08
Δhdp_t	0.39	1.62
$\Delta cons_t$	0.60	3.23
Dummy Q3 2000	0.06	5.37
Dummy Q1 2004	-0.05	-3.96
Adjusted R^2 : 0.69		

Source: Author's calculation.

Finally, we merge these three equations into one model and we forecast estimated the development of household consumption in the period from 1st quarter 2003 to 4th quarter 2005. As the index used for the evaluation of forecasts is higher than in the previous case, we obtain less exact in sample prediction in this case. In the following figures and in table 2.10 we present forecasts starting in the year 2003.

Table 2.10 Nominal household consumption growth (in per cent)

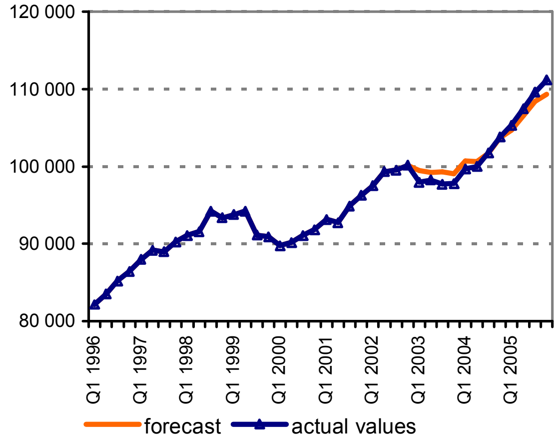
	Q1 2003	Q2 2003	Q3 2003	Q4 2003	Q1 2004	Q2 2004	Q3 2004	Q4 2004	Q1 2005	Q2 2005	Q3 2005	Q4 2005
Actual growth	8.4	6.2	6.1	6.3	12.2	9.9	11.8	11.9	9.4	10.2	10.2	10.1
Forecast (in sample)	10.1	7.3	8.0	7.9	12.1	10.0	10.3	10.5	7.5	8.4	8.7	8.0
Difference (forecast/actual value)	20.1	17.5	31.7	26.0	-1.1	0.6	-12.6	-11.5	-19.7	-17.8	-15.2	-20.7

Source: author's calculation.

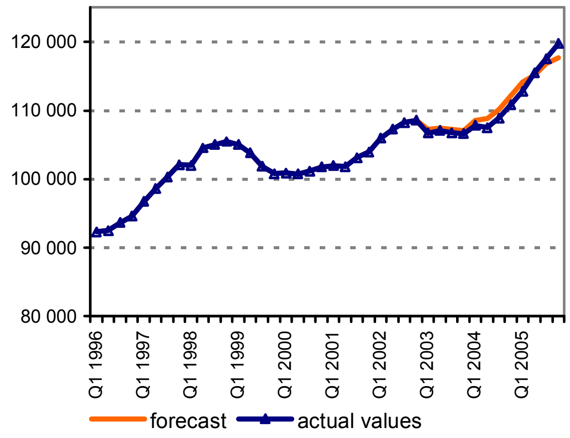
Note: Seasonally unadjusted data is presented.

Figure 11 Dynamic forecast of the 3 equations model (time period 2003:1 2005:4)

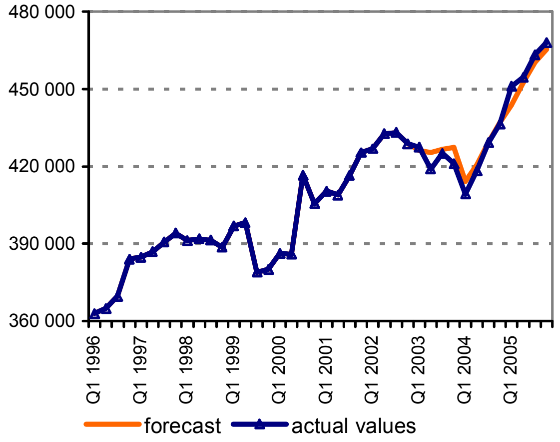
Household consumption



Disposable income



Proxy for wealth (M2+MF)



Measures of predictive ability

	c	y	*a (m2+pf)*
SSF	14860137	13483534	1.86E+08
RMS	1112.8	1060.0	3941.4
RMSP	1.09 %	0.94 %	0.92 %

Note: Time period from 1Q 2003 to 4Q 2005.

Source: Author's calculation.

ARMA model

To offer full information, we introduce also an ARMA model for the household consumption forecasting. We find that the differences in consumption are best characterized with AR(1) vector in combination with MA(3) vectors. In formal words:

$\Delta c_t = \alpha_0 + \rho \Delta c_{t-1} + u_t$, where

$u_t = \varepsilon_t + \theta_1 \varepsilon_{t-1} + \theta_2 \varepsilon_{t-2} + \theta_3 \varepsilon_{t-3}$.

The estimates of this model are presented in the table 2.8. The forecasts created using ARMA model depend only on the previous development of time series of consumption and they are not affected by the real development of economic parameters.

Table 2.11 Estimates of ARMA (1,3)

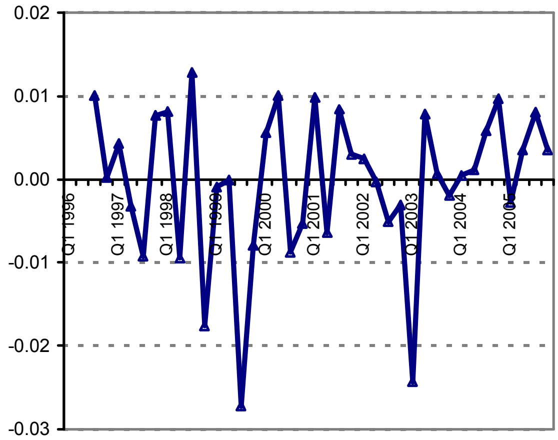
dependent variable: Δc

α_0	0.009
	(2.25)
ρ	-0.58
	(-3.98)
θ_1	0.96
	(11.42)
θ_2	0.93
	(11.12)
θ_3	0.90
	(17.50)
Adjusted R2	0.43

Source: Author's calculation.

Note: t – statistics are in parenthesis.

Figure 2.12 Residuals



Source: Author's calculation.

In sample forecasts of household consumption growth in real prices are presented in the table 2.12.

Table 2.12 Nominal household consumption growth (in per cent)

	Q1 2003	Q2 2003	Q3 2003	Q4 2003	Q1 2004	Q2 2004	Q3 2004	Q4 2004	Q1 2005	Q2 2005	Q3 2005	Q4 2005
Actual growth	8.4	6.2	6.1	6.3	12.2	9.9	11.8	11.9	9.4	10.2	10.2	10.1
Forecast (in sample)	11.1	9.0	10.4	11.4	13.5	11.7	11.1	9.1	7.1	6.0	5.8	6.4
Difference (forecast/actual value)	32.1	44.7	70.2	81.3	11.0	18.2	-5.8	-23.7	-24.0	-40.9	-42.8	-36.8

Source: author's calculation.

Note: Seasonally unadjusted data is presented.

Evaluation of forecasts

As good in sample features of forecast do not guarantee that the model will provide accurate forecasts of the future development, we compare also out of sample prediction for the year 2006 with the actual data (Table 2.13). We find that VEC model overvalue the nominal consumption by approximately 1.3 per cent while other two models undervalue it. Model of the three equations undervalue the forecasts by 1.7 per cent and ARMA model by 0.7 per cent. The mentioned overvaluation of consumption is probably related to the development on labour market. In recent years, we observe a stable rise in employment. The significant part of rise in employment is at present created by increasing the number of people working abroad. Therefore, we use also number of employees calculated using the methodology ESA which does not include people working outside of Slovakia. The estimated coefficient, however, is not significant.

Table 2.13 A comparison of forecasts

	Q1 2006	Q2 2006	Q3 2006	Q4 2006
Actual value	221 462	224 785	235 253	245 659
VEC model				
Forecast (out of sample)	224 597	229 000	237 237	248 109
Difference (in per cent)	1.42	1.88	0.84	1.00
RMSP (in per cent)				1.34
Odhad troch rovníc				
Forecast (out of sample)	221 319	223 636	229 187	236 613
Difference (in per cent)	-0.06	-0.51	-2.58	-3.68
RMSP (in per cent)				2.26
ARMA (1,3)				
Forecast (out of sample)	222 291	226 046	231 497	240 204
Difference (in per cent)	0.37	0.56	-1.60	-2.22
RMSP (in per cent)				1.41

Source: author's calculation.

Note: Seasonally unadjusted data is presented. In current prices.

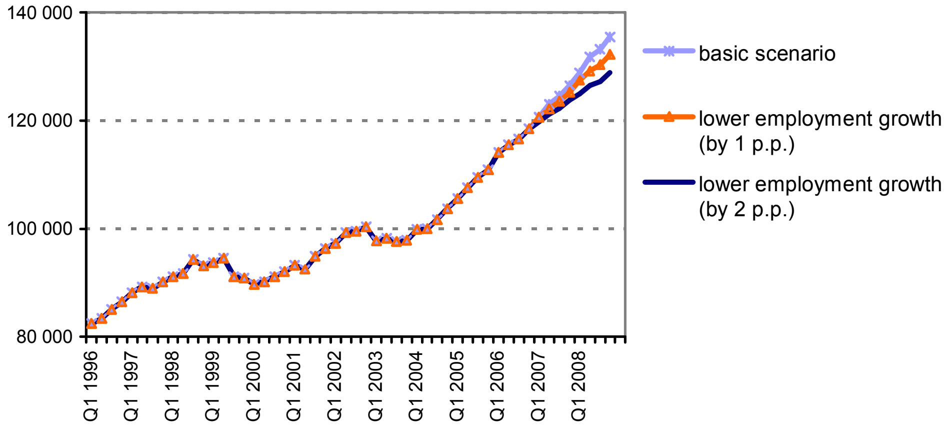
The best model can be chosen on the basis of comparison of the model accuracy. In comparison with other forecasts, the VEC model has the lowest values of RMSP index[23] at both types of prognoses. As a result, we consider the VEC model with exogenous variables (employment and dummy variables) to be the most suitable model of predicting the consumption out of the three presented models. As Slovakia belongs to transforming economies, it is possible that the presented relations will change in the following years and it will be necessary to revaluate the way of estimating the consumption again.

Using the proposed model, we create the forecasts for the years 2007 and 2008. The expected values of the exogenous variable of employment are taken from the mid-term prediction of the National Bank of Slovakia. The chosen model predicts a rise in household consumption by 6.5 - 7 per cent. We also present a sensitivity analysis where we simulate two scenarios. In both of them we expect lower growth of employment (by 1 and 2 p.b.). It turns out that the lowering of the rise of employment by 1 p.b. will show in the lowering of the rise of the final household consumption by 1 to 1.5 p.b. So having a pessimistic

[23] The disadvantage of ARMA model is a very inexact in-sample prognosis with the index RMSP of as much as 2.9 per cent. VEC model had RMSP index of 0.58 per cent at the same in-sample prognosis.

scenario, if the employment will rise slowly (by 2 p.b.) than expected, the real household consumption will rise on average by 4.5 per cent.

Figure 2.13 Forecast sensitivity



Source: author's calculation
Note: In 1995 prices. Seasonally adjusted.

2.4 Conclusions

In this paper we look at the impact of disposable income and financial wealth on the household consumption. We find that financial wealth influences the consumption in Slovakia. We also find the appropriate proxy for financial wealth. It turns out that the most appropriate proxy is a monetary aggregate M2 that represents a significant part of household portfolio extended for assets invested in mutual funds.

Using the Johansen procedure we obtain estimates of elasticities on disposable income (0.93) and on financial wealth (0.34). We assume that high elasticity of disposable income is caused by low share of consumer loans granted to households. As a result, the high share of disposable income is spent on consumption in Slovakia. The result is that the household consumption creates a high portion of disposable income.

We also examine the impact of other relevant variables. We find that the real consumption does not significantly respond to the changes in interest rates. Moreover, the coefficient on consumer confidence indicator is not statistically significant.

The second aim of the presented study is to find an appropriate model for consumption forecasting. We propose and compare three possible ways. The first one is based on the VEC model, the second model consists of three single equations and in the third one we forecast the consumption with ARMA model. Based on a comparison of in sample and out of sample forecasts we prefer to use the VEC model with exogenous variable (employment) for short-time forecasting of household consumption. As Slovakia belongs to the group of transition countries, it is possible that the presented relations will change in the following years and it will be necessary to revaluate the estimates of consumption function.

Appendix

Table 2.14 Model 1 (Vector Error Correction)

1997Q1 2005Q4

t-statistics in [ ]

Cointegrating Eq:	CointEq1		
c_{t-1}	1.00		
y_{t-1}	-0.95		
	[-13.68]		
a_{t-1}	-0.15		
	[-5.28]		
constant	1.20		
Error Correction:	Δc_t	Δy_t	Δa_t
CointEq1	-0.52	0.25	1.02
	[-2.12]	[1.17]	[1.24]
Δc_{t-1}	-0.21	-0.26	-2.19
	[-0.85]	[-1.22]	[-2.69]
Δc_{t-2}	0.76	0.15	-0.12
	[2.73]	[0.63]	[-0.14]
Δc_{t-3}	0.48	-0.09	0.44
	[1.83]	[-0.42]	[0.51]
Δy_{t-1}	0.41	0.38	2.09
	[1.41]	[1.47]	[2.14]
Δy_{t-2}	-0.67	0.19	0.04
	[-2.27]	[0.74]	[0.04]
Δy_{t-3}	-0.63	0.08	-0.91
	[-1.95]	[0.30]	[-0.84]
Δa_{t-1}	0.11	0.16	-0.06
	[2.07]	[3.21]	[-0.36]
Δa_{t-2}	0.07	0.06	0.29
	[1.14]	[1.06]	[1.29]
Δa_{t-3}	0.06	0.04	0.31
	[1.08]	[0.82]	[1.62]
constant	0.001	0.0009	0.009
	[0.96]	[0.48]	[1.45]

R-squared	0.61	0.60	0.45
Adj. R-squared	0.45	0.44	0.23
Sum sq. resids	0.002	0.001	0.02
S.E. equation	0.009	0.008	0.03
F-statistic	3.97	3.85	2.09
Log likelihood	123.15	127.25	80.02
Akaike AIC	-6.23	-6.45	-3.83
Schwarz SC	-5.74	-5.97	-3.35
Mean dependent	0.006	0.006	0.01
S.D. dependent	0.01	0.01	0.03

Determinant resid covariance (dof adj.)	2.56E-12
Determinant resid covariance	8.58E-13
Log likelihood	346.86
Akaike information criterion	-17.27
Schwarz criterion	-15.68

Table 2.15 Model 3 (Vector Error Correction)

1997Q1 2005Q4

t-statistics in [ ]

Cointegrating Eq:	CointEq1		
c_{t-1}	1.00		
y_{t-1}	-0.95		
	[-13.43]		
a_{t-1}	-0.32		
	[-5.18]		
constant	3.81		
Error Correction:	Δc_t	Δy_t	Δa_t
CointEq1	-0.29	0.46	0.25
	[-1.19]	[2.16]	[0.518]
Δc_{t-1}	-0.24	-0.29	-0.37
	[-0.88]	[-1.24]	[-0.69]
Δc_{t-2}	0.53	-0.03	-0.02
	[1.92]	[-0.12]	[-0.04]
Δc_{t-3}	0.51	-0.07	0.55
	[2.15]	[-0.34]	[1.14]
Δy_{t-1}	0.72	0.53	1.06
	[2.67]	[2.25]	[1.97]
Δy_{t-2}	-0.32	0.39	0.19
	[-1.08]	[1.49]	[0.33]
Δy_{t-3}	-0.61	0.11	-1.07
	[-1.86]	[0.39]	[-1.65]
Δa_{t-1}	0.08	0.18	-0.21
	[0.86]	[1.98]	[-1.01]
Δa_{t-2}	0.007	0.06	-0.07
	[0.06]	[0.66]	[-0.35]
Δa_{t-3}	-0.17	-0.04	-0.03
	[-1.71]	[-0.50]	[-0.16]
constant	0.002	0.001	0.004
	[1.18]	[0.86]	[1.03]
R-squared	0.55	0.55	0.29
Adj. R-squared	0.37	0.38	0.01

Sum sq. resids	0.002	0.002	0.01
S.E. equation	0.01	0.008	0.02
F-statistic	3.07	3.14	1.04
Log likelihood	120.46	125.11	95.63
Akaike AIC	-6.08	-6.33	-4.70
Schwarz SC	-5.59	-5.85	-4.21
Mean dependent	0.006	0.006	0.005
S.D. dependent	0.01	0.01	0.02
Determinant resid covariance (dof adj.)		1.18E-12	
Determinant resid covariance		3.95E-13	
Log likelihood		360.83	
Akaike information criterion		-18.04	
Schwarz criterion		-16.46	

Table 2.16 Model 1a (Vector Error Correction)

1997Q1 2005Q4

t-statistics in [ ]

Cointegrating Eq:	CointEq1		
c_{t-1}	1.00		
y_{t-1}	-0.99		
	[-14.85]		
a_{t-1}	-0.12		
	[-4.04]		
constant	1.34		
Error Correction:	Δc_t	Δy_t	Δa_t
CointEq1	-0.54	0.33	0.35
	[-2.84]	[1.76]	[0.45]
Δc_{t-1}	-0.24	-0.32	-2.02
	[-1.36]	[-1.74]	[-2.64]
Δc_{t-2}	0.77	0.06	0.28
	[3.53]	[0.28]	[0.31]
Δc_{t-3}	0.67	-0.05	1.17
	[3.25]	[-0.26]	[1.38]
Δy_{t-1}	0.48	0.39	2.22
	[2.20]	[1.74]	[2.44]
Δy_{t-2}	-0.65	0.31	-0.51
	[-2.84]	[1.37]	[-0.54]
Δy_{t-3}	-0.79	0.12	-1.89
	[-3.08]	[0.46]	[-1.78]
Δa_{t-1}	0.08	0.14	-0.29
	[1.70]	[3.12]	[-1.55]
Δa_{t-2}	0.07	0.06	0.21
	[1.51]	[1.29]	[1.02]
Δa_{t-3}	0.03	0.03	0.19
	[0.85]	[0.77]	[1.08]
constant	0.001	0.001	0.01
	[1.13]	[0.70]	[2.04]
Δzam_t	0.65	0.04	2.61
	[2.67]	[0.17]	[2.60]

Dummy 1Q 2003	-0.02	-0.02	-0.01
	[-3.28]	[-2.78]	[-0.33]
Dummy 1Q 2004	0.02	0.01	-0.01
	[3.03]	[1.90]	[-0.33]
R-squared	0.81	0.75	0.588
Adj. R-squared	0.69	0.60	0.34
Sum sq. resids	0.001	0.001	0.01
S.E. equation	0.007	0.007	0.02
F-statistic	7.22	5.10	2.42
Log likelihood	135.94	135.48	85.050
Akaike AIC	-6.77	-6.74	-3.94
Schwarz SC	-6.15	-6.13	-3.33
Mean dependent	0.006	0.006	0.01
S.D. dependent	0.01	0.01	0.03

Determinant resid covariance (dof adj.)	1.04E-12
Determinant resid covariance	2.37E-13
Log likelihood	370.05
Akaike information criterion	-18.05
Schwarz criterion	-16.07

Table 2.17 Model 3a (Vector Error Correction)

1997Q1 2005Q4

t-statistics in [ ]

Cointegrating Eq:	CointEq1		
c_{t-1}	1.00		
y_{t-1}	-0.95		
	[-15.98]		
a_{t-1}	-0.28		
	[-4.73]		
constant	3.19		
Error Correction:	Δc_t	Δy_t	Δa_t
CointEq1	-0.44	0.42	0.12
	[-2.58]	[2.16]	[0.25]
Δc_{t-1}	-0.22	-0.32	-0.34
	[-1.20]	[-1.54]	[-0.64]
Δc_{t-2}	0.69	-0.01	-0.002
	[3.60]	[-0.06]	[-0.004]
Δc_{t-3}	0.74	0.02	0.70
	[4.48]	[0.15]	[1.44]
Δy_{t-1}	0.74	0.53	1.02
	[4.16]	[2.58]	[1.95]
Δy_{t-2}	-0.46	0.38	0.05
	[-2.26]	[1.64]	[0.08]
Δy_{t-3}	-0.82	0.03	-1.16
	[-3.68]	[0.15]	[-1.77]
Δa_{t-1}	0.04	0.14	-0.28
	[0.59]	[1.83]	[-1.41]
Δa_{t-2}	-0.03	0.03	-0.07
	[-0.55]	[0.43]	[-0.35]
Δa_{t-3}	-0.20	-0.05	-0.13
	[-3.04]	[-0.72]	[-0.69]
constant	0.002	0.002	0.007
	[1.66]	[1.32]	[1.59]
Δzam_t	0.90	0.34	0.87
	[4.27]	[1.40]	[1.41]

Dummy 1Q 2003	-0.02	-0.02	-0.01
	[-3.76]	[-2.90]	[-0.74]
Dummy 1Q 2004	0.02	0.01	-0.03
	[3.05]	[1.43]	[-1.52]
R-squared	0.83	0.71	0.42
Adj. R-squared	0.73	0.54	0.08
Sum sq. resids	0.0009	0.001	0.008
S.E. equation	0.006	0.007	0.02
F-statistic	8.28	4.24	1.25
Log likelihood	137.97	133.06	99.31
Akaike AIC	-6.88	-6.61	-4.739
Schwarz SC	-6.27	-5.99	-4.12
Mean dependent	0.006	0.006	0.005
S.D. dependent	0.01	0.01	0.02
Determinant resid covariance (dof adj.)		4.00E-13	
Determinant resid covariance		9.13E-14	
Log likelihood		387.19	
Akaike information criterion		-19.01	
Schwarz criterion		-17.03	

Table 2.18 Cointegration test (model 1)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	32.316	29.797	0.025
r≤ 1	7.118	15.495	0.564
r≤ 2	0.736	3.841	0.391

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p- value

Table 2.19 Cointegration test (model 2)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	30.036	29.797	0.047
r≤ 1	3.227	15.495	0.956
r≤ 2	0.064	3.841	0.801

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p-hodnota

Table 2.20 Cointegration test (model 3)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	32.432	29.797	0.024
r≤ 1	6.136	15.495	0.680
r≤ 2	0.300	3.841	0.584

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p- value

Table 2.21 Cointegration test (model 4)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	38.353	29.797	0.004
r≤ 1	14.073	15.495	0.081
r≤ 2	0.062	3.841	0.804

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p- value

Table 2.22 Cointegration test (model 1a)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	38.128	29.797	0.004
r≤ 1	8.567	15.495	0.407
r≤ 2	0.179	3.841	0.672

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p- value

Table 2.23 Cointegration test (model 3a)

Number of cointegrating vectors r	Trace-statistics	Critical value (5 per cent)	p-value *
r=0 *	45.339	29.797	0.000
r≤ 1	9.385	15.495	0.331
r≤ 2	1.437	3.841	0.231

Trace test indiates 1 cointegrating vector at 5 per cent level of confidence

*MacKinnon-Haug-Michelis (1999) p-value

3 Wage Flexibility in Slovakia[24]

Wage flexibility is an important concept for monetary policy. It enters into central banks´ thinking about optimum currency areas as well as into its thinking about optimum level of inflation and consequent setting of inflation target.

Knowing the extent of wage flexibility is therefore important in any monetary environment; while having own monetary policy or being a part of larger monetary union. In case domestic monetary policy is present, monetary authorities attempt to set inflation targets with respect to the extent of wage rigidity. If nominal wages are rigid downwards, there may be desirable to accept some inflation to buffer for wage growth especially when its nominal average is close to zero.[25] In case being a part of larger monetary union, other flexible economic policies should be set to compensate for the extent of wage rigidities.

Following extensive literature, we may distinguish two main measures of wage flexibility. The first is the sensitivity of wages to regional unemployment (so called "wage curve") and the other is aversion to wage cuts (so called "downward nominal/real wage rigidity"). In this paper we search for an answer to the second of the two measures on Slovak macro and micro wage data.

Wage setting is in its nature a behavioral process occurring at the level of single economic agents (employees and employers).[26] Wage rigidity in the mind of single economic agent is based on loss aversion, which translates in perceived fairness of wage, which in consequence affects worker's effort (Fehr and Goette, 2000). Therefore, recent studies on wage rigidities prefer using personalized micro data of job-stayers. Since such personalized micro data are not available for Slovakia, our strategy is to start from distribution based aggregate approach and go further into the structure of wage changes in single economic units (enterprises), identify imperfections in measuring transitive economy data with these

[24] This chapter is based on my joint research with Pavel Gertler titled „Downward Wage Rigidities in Slovakia", which was published in Czech Economic Review, vol. 4(1), pages 079-101.

[25] See Groshen and Schweitzer (1999) who closer elaborate on „grease" and „sand" effect of inflation.

[26] These conclusions are confirmed e.g. by findings of Fares and Lemieux (2000) and of Card and Hyslop (1997).

methods and produce the most plausible estimates. Since flexible component in wage (bonuses) is relatively important, the definition of wages as compensations better captures the concept of wage cost of a company.

In contrast to personalized micro-data, using micro-data on the level of enterprises does not allow us to search, where do the rigidities come from. Instead, we will treat rigidities as exogenous, and rather provide for a prudent answer on what is their extent.

A rich list of literature has been devoted recently to the issue of downward wage rigidities. However, to our knowledge only three studies so far used Slovak data to estimate some form of wage rigidity. Blanchflower and Oswald (2000) study uses one year (1995) micro data and finds an elasticity of wages -0.049.[27] Huitfeldt (2001) searches on regional data for effects of unemployment and labor market policies on real wages in Czech Republic and Slovakia in 1992-1998 and finds significantly less wage rigidity in Slovakia than previous study (elasticity under -0.1) as well as compared to Czech republic. Babecký (2008) studies labour market adjustments and also confirms the elasticity below -0.1 with Phillips curve estimates on 1995-1999 aggregate data. He also adds that the relationship cannot be found in the data for Slovakia after 2000 as well as for most of CEE countries. These three studies have looked for some kind of wage flexibility measure through the wage curve and Phillips curve in the past. No studies estimating downward wage rigidity in Slovakia has however taken into account wage change distributions. The aim of this paper is therefore, in the first place, to provide the first estimates of the extent of downward wage rigidity in Slovakia on this basis. In scope of this paper, we understand downward wage rigidity as a share of not realized wage cuts to all wage cuts that should occur in fully flexible environment.

3.1 Concept and literature overview

At the roots of wage rigidities, literature mostly cites Tobin (1972) for his famous claim that "inflation greases the wheels of labour market". He claimed that higher inflation provides for a cushion, in which employer may manipulate wages avoiding nominal wage cuts. A counterforce called "sand effect", i.e. distortionary effects of higher inflation on

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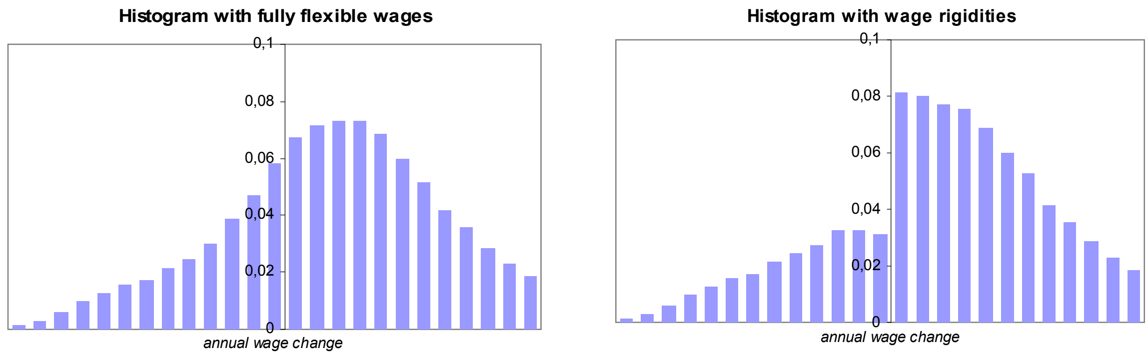
[27] The estimate for Slovakia forms a part of wage curve estimation for central and eastern European countries in the period of 1990-1997. Elasticity equal to -0.1 means that an employee may expect his real wage to decrease by 1% in average if unemployment rate in the region of his workplace grows by 10%, ceteris paribus.

price and wage fluctuations and formation of precise expectations, is then referred to Friedman (1977). A discussion on optimal level of inflation, where both these separated effects cancel out (Groshen and Schweitzer, 1997) used to be regarded as central to monetary policy, because optimum level of inflation provided effective alleviation of wage pressure and thus involving a permanent reduction in unemployment.[28]

Due to level of inflation, wage rigidity has to be considered as a mixed concept of downward nominal and real rigidities (often in literature abbreviated as DNWR/DRWR); however, both being neither alternative nor always simply cumulative concepts. DNWR may become an irrelevant concept in case nominal wage growth is too high (non-effective) or too low (vanish, real rigidities take over). Therefore, downward real wage rigidities are usually more relevant in periods with higher inflation, when nominal growth illusion is being distorted.[29] On the other hand, institutional settings of labor market, especially those indexed by inflation are likely to make real wages more rigid.[30]

Figure 3.1: General shape of distribution without/with wage rigidities



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[28] While discussing optimum level of inflation, findings of Elsby (2006) need to be considered, too. He argues and shows evidence on US and UK micro data that in case of presence of DNWR, besides avoiding wage cuts employers tend to compress also large wage increases in case inflation is high to buffer for future. He therefore concludes that accounting for such weakening of the „grease" effect, optimum level of inflation is to be somewhat lower than previously expected.

[29] Nickell and Quintini (2001) use U.K. NES (New Earnings Survey) data to provide for evidence that spikes at zero nominal wage change are more marked when inflation is low. Besides, they show evidence that since nominal rigidities are focused on zero nominal wage changes, it induces the more distortion in real wage changes the higher is inflation rate (the closer to average nominal wage growth).

[30] Bauer, Bonin and Sunde (2004) conclude that most of the wage rigidity in Germany with central wage bargaining can be attributed to real wage rigidity, which seems to increase with inflation and centralized bargaining outcomes.

Within this paper, we understand downward wage rigidity as a certain number of wage freezes or moderate wage increases that would realize as wage cuts if wages were fully flexible (Figure 3.1). The rationale is that any wage cut causes loss of employee motivation, therefore some moderate wage cuts are too costly for an employer to realize. In such case, it is cheaper for employer to freeze wage or to slightly raise it.

The most common approach used to identify and measure downward wage rigidities in literature in recent decade is the histogram location approach. The problem to be coped with is to compare actual wage change distribution with a notional distribution, which reflects the no-rigidity hypothesis. It is therefore central to define, how notional distribution shall be constructed.

In the method proposed by Kahn (1997), the shape of a distribution of no rigidity hypothesis is assumed to be constant in time. This means that a proportion of observations accumulated in a histogram bar given distance from median should remain constant over time. A presence of DNWR in Kahn's paper is then represented by the extent of misalignment in a relative number of observations in neighboring histogram bars reflecting the position of median. Another method proposed by Lebow et al. (1995) assumes symmetry of a notional wage change distribution. This method comprises simple tail analysis (so called LSW test) of any actual wage change distribution.

Both these assumptions are rather restrictive, though many other factors causing asymmetries and/or non-constant shape of distributions have been found later. For instance Nickell and Quintini (2003) argue that lower inflation supports asymmetry of distribution. They use U.K. Earnings Survey to show that size of spike and step at zero wage growth depends negatively on the rate of inflation (and other statistical parameters of a distribution, too). Lebow et al. (2003) concentrate on other than wage measures that employer may use to compensate for wage cuts, as e.g. cutting social benefits. Consequently they conclude that rigidities may not be seen properly in data. Christofides and Leung (2003) consider the effect of unions´ will to temporarily trade off employment for wage adjustments especially as far as temporary contracts are concerned. Some consented wage freezes may thus infiltrate in data and distort results. Elsby (2006) brings an evidence of leveling off wage changes in time by employers. He argues that especially in case of volatile annual inflation rate and rigid wages, employers tend to restrict wage growth in good times in order to buffer for "low inflation – low wage growth" periods in future. All these effects may have significant effect on symmetry and/or shape of distribution.

Dickens et al. (2006) also begins from a simple symmetry assumption, adjusting however for various possible asymmetries derived from both country data and common cross-country observations in wage distributions. This way they produce one of the most complex and extensive paper so far in this area.

Besides adjustments of Kahn´s or Lebow´s method, some studies use hybrid methods. One of them, by Nickell and Quintini (2003) uses non-parametric estimation of wage change statistical parameters to depict nonlinearities and links wage change distribution to estimated relationships. As the estimation needs to be undertaken on time series, one needs longer time span of data. Yamaguchi (2005) avoids this necessity on Polish data with using more information from wage growth distributions by Kahn-like bar method.

Holden and Wulfsberg (2007) also realize restrictiveness of both assumptions and proposes to construct hypothetical distribution from those actual ones, which they identifies as no-rigidity state. By constructing individual notional distributions from hypothetical distribution adjusting for specific median and variance they conceptually avoid the two restrictive assumptions, too.

Further theoretical and empirical literature reviews on nominal wage rigidity are condensed in e.g. Camba-Mendez (2003). Results of empirical findings are generally in line to conclude with finding an evidence of downward nominal wage rigidity in Europe (Dessy, 2005, Dickens et al., 2006, Knoppik and Beissinger, 2005), its significant cross-country variations (Dickens, 2006 and 2007) and more nominal rigidity in the U.S. compared to Europe (Knoppik and Beissinger, 2005).

3.2 Measuring downward wage rigidities on aggregate data

3.2.1 Data and methodological issues

We use cross-country wage data in sectors of old EU member states extended by 8 new EU member states. The aim of this is to bring in the cross-country factor into the analysis, which allows for comparison of rigidities in wage formation internationally. To do this, we use an unbalanced panel of wage growth data in manufacturing from ILO database. Overall, we have collected 3925 annual wage change observations from 20 countries (EU-25 excluding Malta, Cyprus, Luxemburg, Portugal, Greece and Italy; including Norway), forming 189 country-years in up to 11 year-on-year changes (starting 1996/1995 ending 2006/2005).

Considering our assumptions, we follow Holden and Wulfsberg (2007) approach. In short, the original algorithm goes as follows. All wage change observations in industry j, country i and year t produce a full set S. Then country-years observations with top country-year median wage changes are selected and produce a subset S^{sub}. Observations within S^{sub} are normalized according to specific country-year variances and medians to a vector with overall median of S^{not} equal 0. This is our normalized hypothetical distribution which we suppose reflects full flexibility (no rigidity hypothesis). The shape of this normalized distribution is then laid over all country-year datasets calculating a notional incidence rate for each country-year dataset. This may be interpreted as an incidence rate of wage cuts in case no rigidity hypothesis is valid in the specific country-year set of observations. Empirical incidence rate is normally lower then notional incidence rate, because some wage cuts do not realize. This fraction of unobserved wage cuts (FWCP)[31] represents the extent of rigidity, i.e. share of missing empirical wage cuts with respect to calculated number of notional wage cuts (see technical record in the Appendix 1).

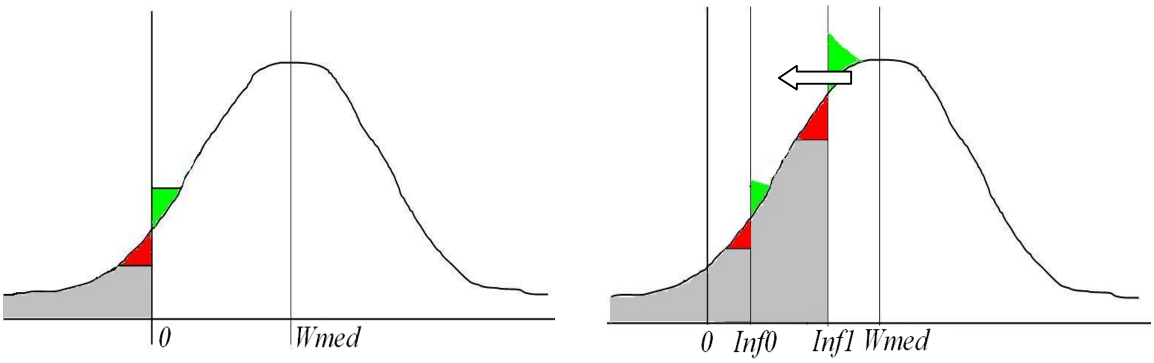
Statistical significance of estimated FWCP is than derived by simulating the probability of a wage cut. This is done by drawing sufficient number of times from binomial distribution (in our case 5000 simulations) of respective number of industries and the probability equal to respective notional incidence rate. We count the number of drawings, where simulated number of negative observations is higher than that of empirical and test it against the null hypothesis of no-rigidity, i.e. empirical incidence rate is equal to notional incidence rate. Such procedure helps us to test the statistical significance of nominal or real rigidity. By conducting the simulation we are able to answer the question whether the simulated extend of the rigidity is significant or not.

In order to derive a non-rigidity hypothesis we need to identify those empirical distributions, where we assume the lowest pressure on zero wage growth. This is solved by picking those empirical distributions, which have the highest nominal and real median wage changes within the sample. While observing that distributions have normal-like shape; the more the vertical *Wmed* line shifts to the right (away from zero wage growth), the higher is the median wage change and the lower shall be the presence of wage rigidities (DNWR - Figure 3.2a and DRWR - Figure 3.2b, adjusted for inflation).

[31] FWCP reflects the proportion of companies, which do not cut the total compensations to companies which would cut the compensations in the absence of rigidity.

We will assume that constructed hypothetical distribution serves as a proxy of no rigidity environment. If such assumption holds, incidence rate of wage cuts yielding from this hypothetical distribution would equal to empirical incidence rate.

Figure 3.2a, 3.2b: Downward nominal and real wage rigidities on histogram



Note: portion of wage change observations below zero nominal/real wage growth (red portion of grey field) transforms into wage freezes or slight wage increases (green) forming a spike at zero growth and/or step like shape of distribution around zero growth.

In our sample, maximum number of industries per country-year is 23, minimum is 12. Overall, we observe 235 nominal wage cuts, i.e. incidence rate of 6% of all observations. These wage cuts are distributed unevenly, all of them within 100 country-years; other 89 country-years do not include a single industry, with year-on-year wage decline in nominal terms. Further statistics of the sample is presented in the following table.

Table 3.1: Statistics of the data sample

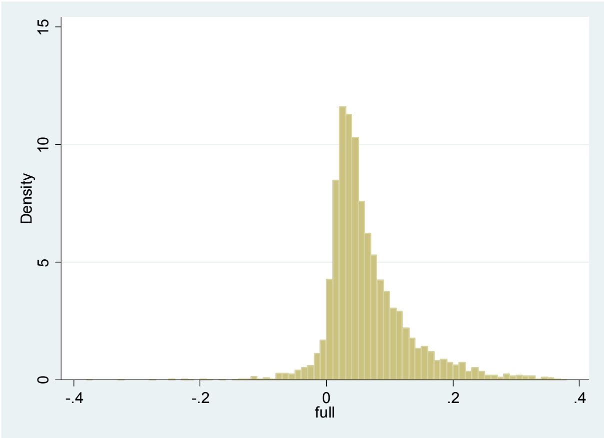
	Total sample	1996-2000	2001-2006	new EU	V4 countries	Baltics	Slovakia	EU-15
number of observations	3925	1616	2309	1817	962	612	236	20
number of countryyears	189	78	111	85	43	31	11	[illegible]
inflation	0.040	0.057	0.028	0.064	0.069	0.057	0.068	0.0
Statistics nominal wage changes								
number of nominal wage cuts	235	95	140	93	23	67	2	1
number of country-years with no nominal wage cuts	89	40	49	49	29	12	9	[illegible]
standard deviation of nominal wage changes	0.073	0.090	0.054	0.088	0.076	0.116	0.061	0.0
P75-P35 - nominal wage changes	0.056	0.084	0.039	0.068	0.072	0.096	0.036	0.0
median of nominal wage changes	0.048	0.065	0.042	0.088	0.088	0.090	0.091	0.0
incidence rate	6.0%	5.9%	6.1%	5.1%	2.4%	10.9%	0.8%	7.1
share of countryyears with no cuts	47.1%	51.3%	44.1%	57.6%	67.4%	38.7%	81.8%	32.6
Statistics real wage changes								
number of real wage cuts	926	382	544	423	229	159	70	5
number of country-years with no real wage cuts	21	9	12	7	3	4	0	
standard deviation of real wage changes	0.061	0.074	0.049	0.079	0.061	0.111	0.062	0.0
P75-P35 - real wage changes	0.034	0.038	0.032	0.047	0.041	0.071	0.044	0.0
median of real wage changes	0.019	0.021	0.018	0.030	0.028	0.044	0.024	0.0
incidence rate	23.6%	23.6%	23.6%	23.3%	23.8%	26.0%	29.7%	25.0
share of countryyears with no cuts	11.1%	11.5%	10.8%	8.2%	7.0%	12.9%	0.0%	7.4

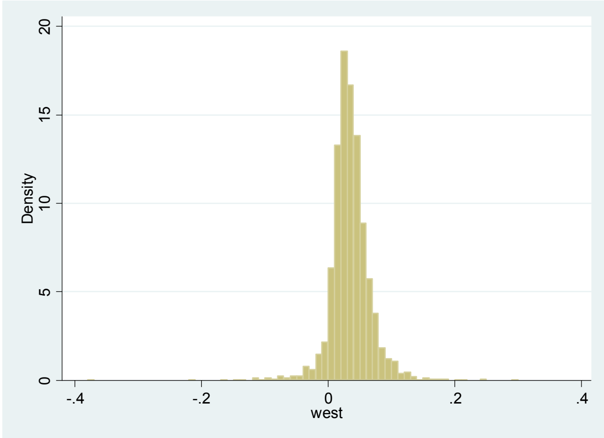
Source: authors' calculation.

3.2.2 Examining effects of full sample heterogeneity

The fact that our data come from both developed and transition economies results in significantly different statistics for these two groups.[32]

Figure 3.3a, 3.3b: Histogram of nominal wage changes in full sample (left), EU-15 economies and Norway (right).





Source: authors' calculation.

Full sample of raw wage changes data is far more skewed to the right with much lower kurtosis, because of higher nominal wage growth in transition countries mainly due to convergence process.[33] Different statistics of the two subsamples of raw wage changes suggest that we should examine, whether some effect of this disparity is transferred into other relationships.

To illustrate features of the two subsamples, we searched for the sensitivity of incidence rate to median wage growth. In average in our data sample, one percentage point shift in median real wage growth to the left translates into 4.3 to 6.3 percentage points more real wage cuts (causing higher incidence rate of wage cuts)[34]. The same size of shift in nominal terms translates into 0.9 to 2.1 percentage points more nominal wage cuts. These findings confirm higher sensitivity of incidence rate in real terms due to smaller distance of wage changes from the level of inflation than is their distance from nominal zero growth (see Figure 3.2a/3.2b). Looking at separate subsamples however, we produce varying

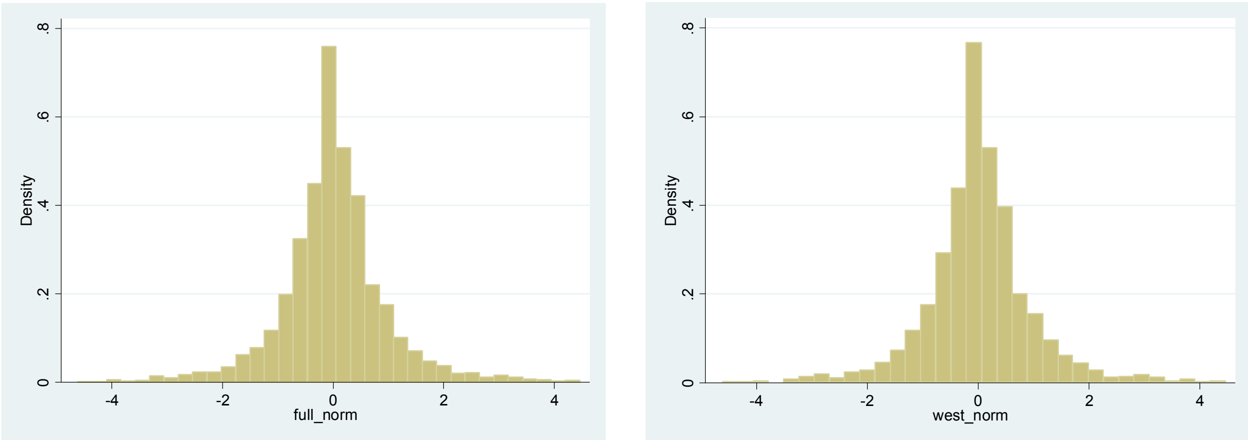
[32] Further in the text to be referred to as EU-15 for developed economies and EU-10 for transition economies.

[33] Skewness of the full sample is 1.19 compared to 0.26 of EU-15 countries; kurtosis of the full sample is 15.6, compared to 28.1 of EU-15. Mean nominal wages of EU-15 is 3 percentage points lower than of the full sample, resp. 1.5 percentage points in terms of median nominal wages.

[34] Underlying relationship is non-linear. Inspired by Nickell and Quintini (2001) we regress incidence rate of wage cuts on respective median wage change and its square and standard deviation.

coefficients. We observe that sensitivity of incidence rate to nominal wage growth for EU-15 countries increase to 2.0 to 5.5 percentage points and to real wage growth to 11.9 to 27.7 percentage points. To compare, coefficients for EU-10, both in nominal and real terms remain roughly the same as in the full sample.

Figure 3.4a, 3.4b: Histogram of normalised nominal wage changes in full sample (left), EU-15 economies and Norway (right).



Source: authors' calculation.

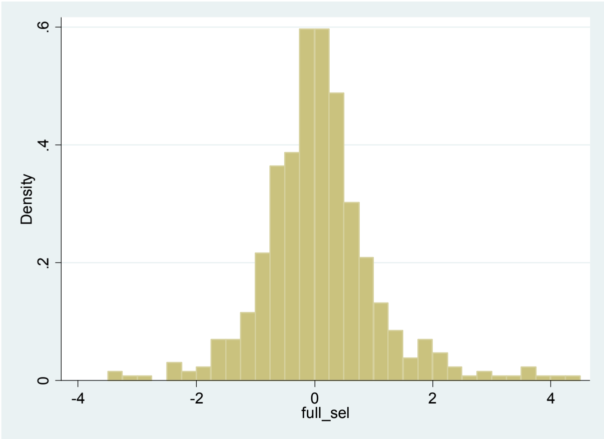
If observing these data after normalization, varying features of geographic subsamples do not cause much of a problem. Distribution of full sample of countries easily passes the test of equality of distributions[35] with the one of EU-15 subsample.

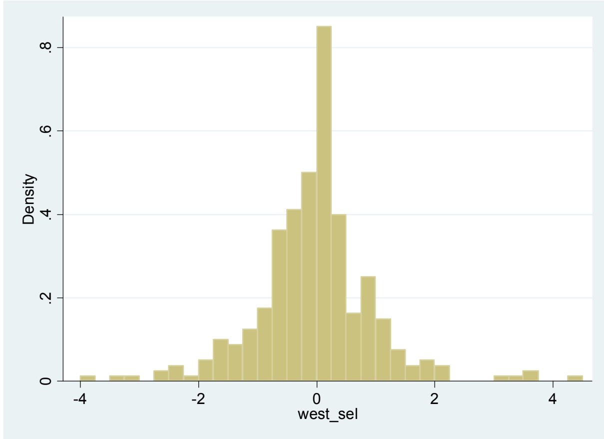
Distinct character of raw data for the two geographic subsamples then does not represent any problem if normalized and if resulting notional incidence rates and their relationship to empirical incidence rates are not distorted due to this disparity.

Perhaps more sensitive issue is a selection procedure, which identifies such country-years, that shall be assumed to represent non-rigidity environment. Holden and Wulfsberg (2007) suggest populating hypothetical distribution with those empirical distributions, where median wage growths (both nominal and real) qualify in their respective upper quartiles. As we will first reproduce this approach on extended number of more unlike countries, such qualifying criteria will depend on statistical parameters of raw wage changes (not-normalized). Having geographical subgroups with such median wage changes in one pool, criteria will always be met only by the EU-10 country-years, which posses higher median values. Resulting distributions of "non-rigidity assumed" normalized wage changes of full sample compared to those of EU-15 subgroup already yields some visible differences.

[35] Tested with two-sample Kolmogorov-Smirnov test of equality.

Figure 3.5a, 3.5b: Histogram of nominal wage change observations qualifying in the upper quartile selection in full sample (left), EU-15 economies and Norway (right).





Source: authors' calculation.

While controlling for medians, mean is located more to the left in EU-15 subgroup, moreover distribution is less positively skewed. Nevertheless, testing for equality (by two-sample KS test) does not rule out that the two distributions are alike.

However, we need to pay more attention to the harmony of the two steps of methodology: construction of hypothetical distribution and derivation of notional distributions. Since quartile selection in full sample is being drawn from right fat tail of the distribution (fig. 3a), it comprises only of EU-10 country-years. Incidence rate (the share of negative observation) in raw wage change data of EU-15 and full sample is about the same level, but the two sets differ largely in symmetry (full sample data being positively skewed, while EU-15 more or less symmetric - Figure 3.3a/3.3b). Right tail observations drawn from the full sample then represent very much of an outlier relative to EU-15 set of observations. Even though normalization through median and variance absorbs much of the noise, notional incidence rates of EU-15 country-years are for this reason subject to downward shift by 1 to 2.5 percentage points (difference of notional incidence rates in Table 3.2 and Table 3.3). We may therefore conclude that resulting FWCP of EU-15 country-years are pushed downwards (by 25% in average in proposed composition) if they are calculated in full sample with EU-10.

Having no other attributes that could possibly serve as a proxy for non-rigidity environment, one may think of some manipulation of qualification criteria. If we do that and narrow the criteria from upper quartiles to deciles, hypothetical distribution remains with fewer selected observations of even more outlier data. The above described effect is then even stronger. Besides, hypothetical distribution with fewer observations produces larger risk of non-normality, further hurting reliability of notional incidence rates and consequent FWCP. For further calculation we will therefore stick to the selection of at least top quartile observations.

No matter how is the criterion set, we are able to estimate the extent of possible downward shift. We shall not yet therefore disqualify the approach in effort to estimate downward wage rigidities in our full sample; just interpret it with caution.

3.2.3 Results and their robustness

When a full sample of both transition and developed economies was used, only EU-10 countries´ country-year data had classified into the hypothetical distribution. Greater variance and median changes of these country-years then reflected into the distribution, giving distorted information to notional distributions of developed countries (EU-15). Further to this, low empirical incidence rates mainly in EU-10, distort the calculated fraction of wage cuts prevented -FWCP).

Table 3.2: Downward wage rigidity of full sample

Countries	No.of years	All Obs	q_emp	q_not	q_sim	fwcp sim	p-val
Austria	7	161	0.0186	0.0601	0.0601	0.6897	0.0106
Belgium	5	114	0.1316	0.1953	0.1955	0.3269	0.0514
Denmark	10	230	0.1043	0.1078	0.1077	0.0309	0.4908
France	9	207	0.0193	0.0325	0.0325	0.4056	0.1960
Ireland	10	120	0.0167	0.0155	0.0157	-0.0632	0.7130
Hungary	11	253	0.0237	0.0164	0.0165	-0.4412	0.8772
Finland	10	228	0.1228	0.1320	0.1324	0.0725	0.3760
Estonia	10	172	0.0407	0.0357	0.0351	-0.1586	0.7432
Latvia	11	232	0.1336	0.1667	0.1660	0.1949	0.0792
Lithuania	10	208	0.1394	0.1486	0.1495	0.0674	0.3836
Netherlands	10	190	0.0421	0.0552	0.0553	0.2389	0.2644
Norway	9	102	0.0000	0.0006	0.0007	1.0000	0.9354
Poland	11	253	0.0435	0.0449	0.0453	0.0401	0.5204
Slovakia	11	236	0.0085	0.0442	0.0423	0.7995	0.0022
Slovenia	11	243	0.0123	0.0072	0.0072	-0.7233	0.9022
Spain	7	161	0.1056	0.0731	0.0733	-0.4399	0.9582
U.K.	9	207	0.0918	0.1049	0.1045	0.1215	0.3158
Sweden	8	159	0.0377	0.0429	0.0433	0.1284	0.4644
Germany	10	229	0.0699	0.0674	0.0668	-0.0462	0.6478
Czecb rep.	10	220	0.0182	0.0167	0.0165	-0.1014	0.7032
	189	**3925**	**0.0599**	**0.0654**	**0.0684**	**0.1242**	**0.0124**

Note: AllObs – number of all observations, q_emp – empirical incidence rate, q_not – notional incidence rate, q_sim – simulated notional incidence rate, fwcp_sim – simulated fraction of wage cuts prevented, p-val – probability of significance.

Source: authors' calculation.

Comparing nominal wage rigidities across countries from the full sample should be carried out with caution. Calculating the same for EU-15 sample separately yields fully comparative results of wage rigidities. Comparing notional incidence rates and fractions of wage cuts prevented for relevant countries between the Table 3.3 and the Table 3.2 we may observe differences, which occur when distributions with higher wage changes enter into the sample.

In standard situation higher FWCP means more rigidity, FWCP approaching zero means more downward wage flexibility. Negative FWCP values are present in situations, where calculated notional incidence rate of wage cuts is lower than measured empirical incidence rate of wage cuts. Such situation may therefore also be considered as wage flexibility. All these findings apply only in case the two incidence rates are significantly distant (p-value) and therefore disturbances are eliminated.

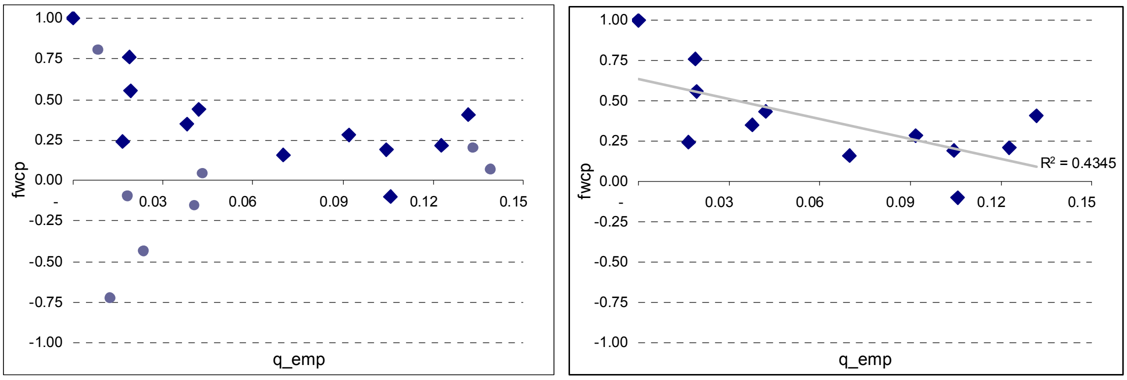
Table 3.3: Downward wage rigidity- old EU-15

Countries	No.of Years	All Obs	q_emp	q_not	q_sim	fwcp sim	p-val
Austria	7	161	0.0186	0.0763	0.0766	0.7567	0.0006
Belgium	5	114	0.1316	0.2200	0.2213	0.4053	0.0104
Denmark	10	230	0.1043	0.1294	0.1294	0.1937	0.1400
France	9	207	0.0193	0.0434	0.0434	0.5547	0.0548
Ireland	10	120	0.0167	0.0222	0.0219	0.2386	0.5140
Finland	10	228	0.1228	0.1550	0.1556	0.2110	0.0964
Netherlands	10	190	0.0421	0.0750	0.0749	0.4375	0.0448
Norway	9	102	0.0000	0.0017	0.0017	1.0000	0.8378
Spain	7	161	0.1056	0.0960	0.0963	-0.0970	0.7102
U.K.	9	207	0.0918	0.1281	0.1276	0.2805	0.0682
Sweden	8	159	0.0377	0.0570	0.0578	0.3474	0.1772
Germany	10	229	0.0699	0.0841	0.0833	0.1609	0.2614
	104	**2108**	**0.0674**	**0.0863**	**0.0944**	**0.2863**	**0.0000**

Source: authors' calculation.

As it arises from (3.5) in appendix, fraction of wage cuts prevented should be negatively sloped relative to empirical incidence rate. This is true in EU-15 sample, however not so in the full sample (See Figure 3.6a, 3.6b).

Figure 3.6a, 3.6b: Empirical incidence rate and FWCP in full sample (left), EU-15 economies and Norway (right).



Source: authors' calculation.

Points in the graph 6a do not visually follow the logic of the equation as they do in 6b, because of high insignificance (p-val) of FWCP in EU-10.

Low empirical incidence rate is caused by high median wage change (Table 3.1). Small number of industries per country-year reflects in discontinuous values of incidence rate with relatively high increments. If incidence rates are close to zero, fraction of wage cuts prevented becomes extremely sensitive to random disturbances, e.g. between no-wage-cut-recorded situation and one–wage-cut-recorded situation. Consequently, the countries with low aggregate empirical incidence rate tend to be more exposed to sensitivity to random disturbances in individual country-years (prone to distortions) and therefore do yield more variable FWCP values. Resulting FWCP of countries with empirical incidence rate in immediate distance from zero (empirical evidence suggests less than 0.02, which in our data translates to at most 3 overall negative observations) does not testify properly about the wage rigidity. Results on wage rigidity in such countries (Slovakia, Slovenia, Austria, Ireland, Norway) are then difficult to interpret.

Thus we may conclude that if calculating wage rigidity for developed and transition economies together, we face two possible sources of distortion. First, coming from construction of non-rigidity hypothesis of two distinct sets of data and second from too few wage cut observations, which origin partly in higher median wage changes.[36]

[36] Besides Slovakia and Slovenia, Ireland and Norway may be considered so due to its excess growth relative to EU average in the time observed.

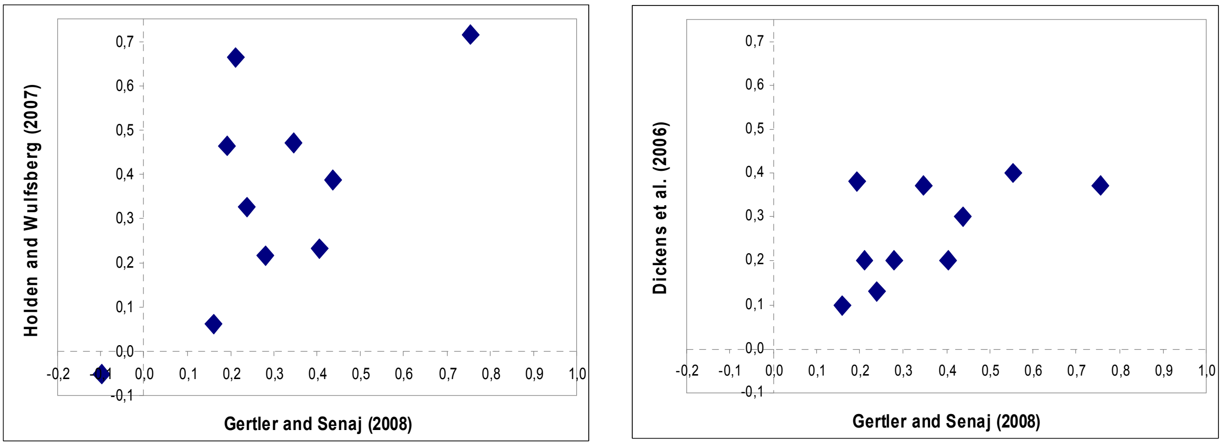
3.2.4 Results - interpretation

From the above reported tables we may confirm some downward nominal wage rigidity (29%) in the EU-15 sample. Based on 5000 simulations, we observe nominal wages being significantly rigid downwards especially in Austria, France, the Netherlands and Belgium, where over 40% of wage cuts are being prevented.[37] For all the other countries of the EU-15 sample we have not found significant wage rigidities, even though we found over 20% downward wage rigidities in U.K. and Finland on 10% significance level. These findings are generally in line with other results from cross-country studies., whose downward nominal wag rigidity estimate of the full sample reaches 26%.[38]

Results for downward nominal wage rigidities for separate EU-15 sample may be confronted with past evidence of identical approach of Holden and Wulfsberg (2007) and of Dickens et al. (2006).

For EU-10 countries we may confirm more downward wage flexibility than in EU-15, while most of the significant FWCP values are closer to zero. However, for most of the EU-10 countries in the sample we could not measure any wage rigidities (being insignificant). From the data we have, we cannot conclude for Slovakia. This is because of very few negative observations and therefore being exposed to random disturbances.

Figure 3.7: Our country results of FWCP compared to country results of Holden and Wulfsberg (2007) and Dickens et al. (2006).



Source: authors' calculation.

[37] We do not consider Norway, where there is no empirical wage cut observed, therefore FWCP is always equal to one no matter all other parameters. Testing however proves the high value of Norwegian FWCP being insignificant.

[38] They include some extra OECD countries and uses data from longer period further in history (1973-1999).

The ability to compare downward wage rigidities between old and new EU member states was the supporting idea for choosing the histogram location approach. However, now we see that due to structural differences in our data, any effort to estimate the full sample together leads us to incomplete information.

3.2.5 Reasons for turning to microdata

We have identified two sources of possible distortion to the output data while applying histogram location approach of Holden and Wulfsberg (2007) on mixed sample of old and new EU member states wage growth data.

The first is an incomparability of resulting FWCP in transition economies and developed economies. It origins in heterogeneity of median wage changes, rates of inflation and in variance of country-year observations in the two subsets. Much of this heterogeneity is absorbed by normalisation, however selection criteria to construct non-rigidity hypothesis inclines to the EU-10 country-years. Notional incidence rates are then shifted downwards and consequently are the extent of wage rigidity upwards.

The second is the deficiency coming from too few observations in partial distributions (country-year). Having only 20 observations (industries) in one country-year may be a source of noise, which may be crucial if empirical incidence rate is too close to zero. This is particularly true in some countries, which makes interpretation of wage rigidity for these countries more complicated.

The last, but not less important is the difficulty in interpreting resulting FWCP. Ideally, we would like to say that some share of wages was prevented from dropping over the year. Since our single observation is an entire industry/sector, we may only say that certain percentage of average sectoral wage cuts over the year was prevented.

Knowing that we lose much of within-sector wage information and being unable to reliably estimate wage flexibility for Slovak republic, we turn our attention to microdata. As individual chained wage data are not available for Slovakia (neither in most of the transition countries), we will use the micro-data on company level. Results of aggregate data will serve as a reference and a useful starting point to compare all the next results with.

3.3 Measuring wage rigidity on the company level

3.3.1 Data

Departing from the findings above, we put the emphasis on the analysis of the company level micro-data. To our knowledge, company level microdata have been used in histogram location approach in two studies so far. In Lebow et. el. (2003) wages are defined as hourly costs of wages and benefits in an establishment. Their data source is the Bureau of Labor Statistics' employment cost index. Likewise, Brzoza-Brzezina and Socha (2007) employed enterprise level data from a survey of medium sized and large enterprises in Poland. Besides other findings, both of these papers provide evidence that the total compensations are less affected by the downward nominal wage rigidity than base wages alone. Since we are use similarly compensation-type wage definition, we shall account for larger flexibility from margins of adjustment in flexible components of wage.

We employ similar type of data source as Brzoza-Brzezina and Socha (2007). Since we cannot track individual wages over time in Slovak data we find business surveys conducted annually by the Statistical Office of the SR as the most appropriate data sources for this type of analysis in Slovak environment. Particularly, three surveys[39] were merged in order to obtain as representative sample as possible. We use total compensations in a company. Although small businesses (up to 19 employees) are not fully represented in the database, (this is one of the drawbacks of our data source) medium (from 20 to 99 employees) and large companies (with more than 100 employees) are surveyed exhaustively. The database used covers about half of the employees in the production sector of the economy. The following table compares data for the economy as a whole and the sample used.

Table 3.4: Comparison of the data sample and the Slovak economy.

	2007	2006	2005	2004	2003	2002	200
data set							
Number of enterprises	5 498	5 494	5 039	4 932	5 138	4 904	4 81
Respective number of employees	834 749	849 470	732 986	749 790	790 487	735 650	774 87
slovak economy (production sector)							
Number of employees	1 766 541	1 712 702	1 668 034	1 621 704	1 616 513	1 608 622	1 607 55

Source: Statistical Office of the SR, (SO SR), authors' calculation.

[39] e.g. Annual questionnaire on business statistics (ROC 1-01), Annual questionnaire in banking and non-banking financial institutions (PEN P 5-01), Annual questionnaire in insurance (POI P 5-01).

We consider both full time and part time employees. The central variable we use is the change of average hourly compensation (in both nominal and real terms) in the company.[40] Further, we filter the database to eliminate an impact of assumed error inputs, which originate mainly from incorrectly filled in questionnaires[41].

The dataset covers the period from 2000 to 2007. Due to the methodological changes in the surveys, the years before 2000 are not considered. Selected time period includes years with lower (2.8 %) as well as higher (8.5 %) level of inflation. The difference between highest and lowest inflation rate is almost 5.7 what guarantees that the distributions of changes in compensations are different across the sample.[42] The basic statistical properties of the analyzed data sample are presented in Table 3.5.

Table 3.5: Statistical properties of the changes in compensations and basic macro indicators

	2007	2006	2005	2004	2003	2002	200
Statistics of changes in nominal total compensation							
Median	0.084	0.078	0.063	0.054	0.063	0.101	0.06
Mean	0.084	0.078	0.063	0.062	0.060	0.101	0.06
Standard deviation	0.136	0.135	0.132	0.137	0.134	0.138	0.13
Statistics of changes in real total compensation							
Median	0.056	0.033	0.036	-0.021	-0.022	0.068	-0.00
Mean	0.056	0.033	0.036	-0.013	-0.025	0.068	-0.00
Macro indicators							
Unemployment rate [%]	11.0	13.3	16.2	18.1	17.4	18.5	19.
Employment growth [%]	2.4	3.8	2.1	0.3	1.8	0.2	1.
Average wage growth [%]	7.2	8.0	9.2	10.2	6.3	9.3	8.
Inflation rate [%]	2.8	4.5	2.7	7.5	8.5	3.3	7.

Source: SO SR, authors' calculation.

An interesting difference between industry and company level data can be seen in the Figure 3.8, which shows the distributions of the annual changes in total compensations. Although Slovak industry level data (used in the previous part) displays hardly any wage cuts during the whole sample, almost 30% of observed companies cut their hourly

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[40] $\Delta wage_{t/t-1} = wage_t / wage_{t-1} - 1$. For wages we use the definition of compensations, i.e. wages including bonuses and premiums since partial components of wages are not available from this survey. The total amount of wage costs were divided by the total amount of hours worked. Although, both numbers are reported by companies, such definition may lead to measurement errors.

[41] Annual change of more than 50% to the level of compensations is considered as incorrect input in any of the two years and such observation is therefore eliminated. Observations with missing values were eliminated, too.

[42] Kramarz (2001) claims that wage change distributions in years of high inflation strongly differ from those observed in years of low inflation.

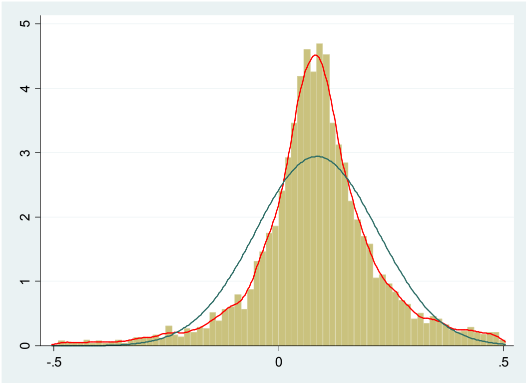
compensations.[43] Considerable different results compared to sectoral analysis are most likely a mixed effect of above mentioned aggregation issue, flexibility of employers in handling the number of hours worked or by the fact that company level micro-data includes over 60% of other observations than the ones from manufacturing.

Another interesting issue is that despite strong average wage growth of 6 % to 10 % during the years 2000 and 2007, almost 30% of analyzed enterprises cut their hourly compensations in each year. This paradox may be explained by at least the following three reasons. Firstly, changes in the composition of workforce may have changed the average compensation even if the wage rates stayed on the same level. Secondly, changes in the number of hours worked may have modified the average hourly compensation even if hourly wage remained the same. And finally, cutting bonuses in aiming to decrease total costs of the company could also lower total compensations. It has been shown by Babecky et. al (2008), that changes in bonuses, non-pay benefits and slowing down promotions belongs to potential margins used by companies to reduce labor costs. They also present survey results on the particular case of the Czech Republic that 31 % of companies prefer to reduce bonuses, 9 % prefer cheaper hires, 9 % choose early retirements and 50 % of the companies use other labor cost reduction strategy.

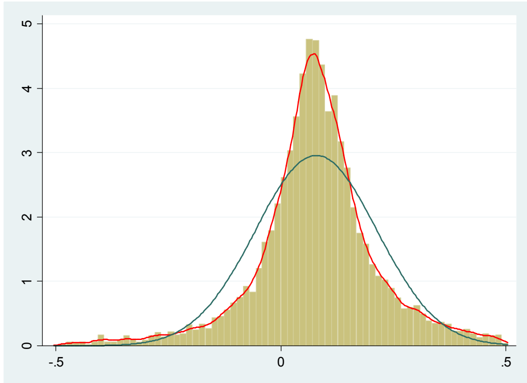
[43] Blinder and Choi (1990) discovered that the money wage cuts were more common in the US than they had imagined even they analyzed a time period characterized by low unemployment.

Figure 3.8: Distribution of changes of hourly compensations. Kernel vs. normal density functions.

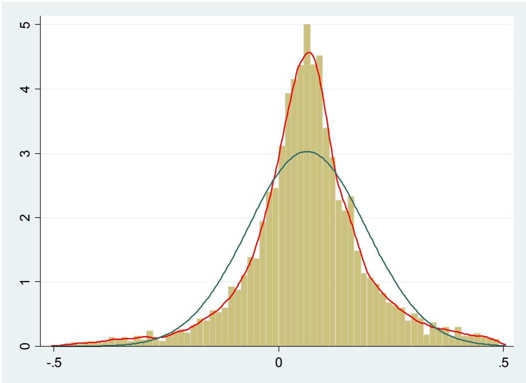
2007/2006



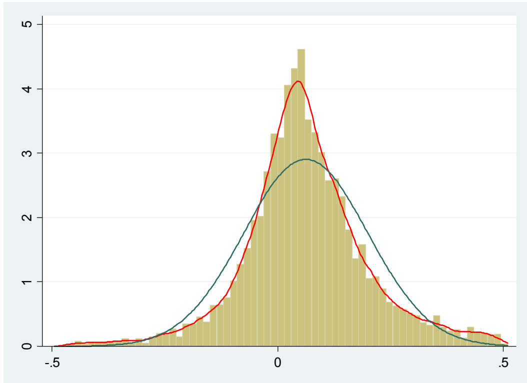
2006/2005



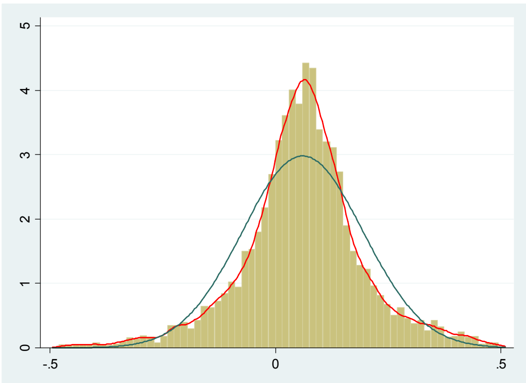
2005/2004



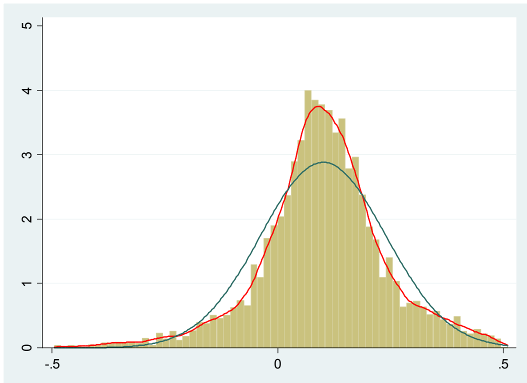
2004/2003



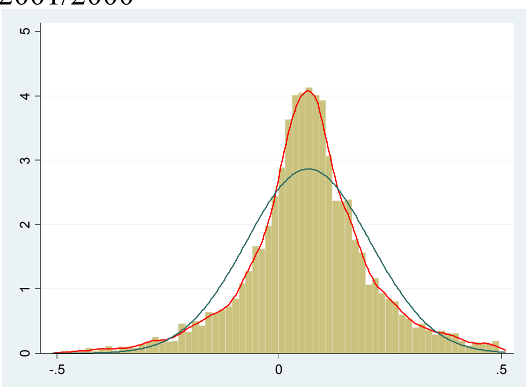
2003/2002



2002/2001



2001/2000



Source: authors' calculation.

3.3.2 Results

In order to apply the original approach proposed by Holden and Wulfsberg (2007) to company level data, we slightly modify their method of choosing the hypothetical (underlying) distribution. We also assume no rigidity in the hypothetical distribution, but here we are constrained by shorter time period. The analyzed data sample consists of seven years, thus we pick only one year (instead of bulk of country years as in chapter 3.2) out of our sample with highest median of nominal and real growth of the total hourly compensations. According to our data, there is no doubt for choosing the year 2002.[44] Nevertheless, there is still possibility, that wage rigidity was present also in 2002 data. If this is the case, the presented figures stand for the lower bound of the actual extent of rigidity.

An empirical investigation shows interesting results. Table 3.6 presents the outcomes of analysis of nominal rigidity in total hourly compensations. In the early years of the sample we did not find a presence of rigidity. Notional incidence rate (*q_not*) significantly exceeds empirical incidence rate only after 2005. Consequently, the fraction of wage cuts prevented rises from about 5 % in 2005 to almost 10 % in 2007. The estimated FWCP are statistically significant. Thus we can conclude that at least 5 % out of those companies, which would cut compensations in the absence of rigidity, are affected by downward nominal wage rigidity (in 2005).[45] Another important finding is that the degree of rigidity tends to slightly increase in recent years. For the sake of simplicity we calculated shares of the companies affected by nominal rigidity and they are reported in column labeled as nominal wage rigidity (nwr). Nwr ranges from 1.5 % to 2.2 %, which means that at least 1.5 % of companies were affected by wage rigidity in 2005.[46]

Our results are in line with those reported for Poland. Brzoza-Brzezina and Socha (2007) concluded that the extent of rigidity at the enterprise level was relatively small during the period 1996 – 2005.

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[44] Further we follow the algorithm described in the Appendix 1.

[45] Since the obtained results could be affected by adjustments in the company structure (such as by substituting expensive employees by cheaper ones), Brzoza-Brzezina and Socha (2007) suggest to treat the results as the lower bound of the true DNWR at enterprise level.

[46] It is important to stress that the results may be partly influenced by the business cycle. During the period studied, Slovakia recorded strong economic growth (employment growth can be found in Table 5).

Table 3.6: Nominal wage rigidity – Slovakia

Years	All Obs	q_emp	q_not	q_sim	fwcp sim	nwr	p-val
2007/2006	5498	0.194	0.214	0.214	**0.096**	0.021	0.000**
2006/2005	5494	0.207	0.229	0.229	**0.094**	0.022	0.000**
2005/2004	5039	0.248	0.263	0.263	**0.058**	0.015	0.006**
2004/2003	4932	0.288	0.299	0.299	**0.036**	0.011	0.051
2003/2002	5138	0.268	0.268	0.268	**0.000**	0.000	0.504
2002/2001	4904	0.185					
2001/2000	4812	0.257	0.266	0.266	**0.033**	0.009	0.085

**Note: DNWR are statistically significant at 1% level of significance.
Note: AllObs – number of all observations, q_emp – empirical incidence rate, q_not – notional incidence rate, q_sim – simulated notional incidence rate, fwcp_sim – simulated fraction of wage cuts prevented, nwr – share of companies affected by nominal rigidity, p-val – p value.
Source: authors' calculation.

An interesting question arises about the impact of detected rigidity in hourly compensations on the labor market, particularly on wage growth (and consequently on inflation). The estimated impact of downward nominal wage rigidity on wage growth is relatively low and can be considered negligible. For instance, in 2006 (FWCP equals to 9.4 %) downward nominal wage rigidity caused additional costs to employers in amount of 296 million Sk (9.8 mil. EUR). If we translate this to annual wage dynamics, this amounts to 0.14 percentage points of the wage growth if compared to fully flexible environment.

Applying the same methodology on inflation adjusted data; the extent of the downward real wage rigidity can be analyzed.

Table 3.7: Real wage rigidity - Slovakia

Years	All Obs	q_emp	q_not	q_sim	fwcp sim	rwr	p-val
2007/2006	5498	0.268	0.289	0.289	**0.072**	0.021	0.000**
2006/2005	5494	0.347	0.371	0.371	**0.063**	0.024	0.001**
2005/2004	5039	0.343	0.353	0.353	**0.028**	0.010	0.080
2004/2003	4932	0.584	0.578	0.578	**-0.011**	NA	0.811
2003/2002	5138	0.598	0.585	0.585	**-0.022**	NA	0.969
2002/2001	4904	0.259					
2001/2000	4812	0.530	0.525	0.525	**-0.008**	NA	0.732

**Note: DRWR are statistically significant at 1% level of significance.
Note: AllObs – number of all observations, q_emp – empirical incidence rate, q_not – notional incidence rate, q_sim – simulated notional incidence rate, fwcp_sim – simulated fraction of wage cuts prevented, rwr – share of companies affected by real rigidity, p-val – p value.
Source: authors' calculation.

It turns out that real wage changes are affected by real wage rigidity only in the last two years, FWCP grows from 6 % in 2006 to 7 % in 2007. The extent of real rigidity measured as a share of companies affected by real rigidity (column labeled as rwr in Table 3.7) is almost comparable to the share of companies affected by nominal wage rigidity. It should be noted that applying wide definition of wage (including bonuses) makes it easier for employer to adjust pays in any of the years; therefore level of reported rigidities represents its minimum bound.

The overall wage rigidity may not correspond to those in different segments of corporate sector. Next, we therefore measure the degree of rigidity in different subgroups classified by company size and sector of economic activity (according to primary NACE classification). Firstly, we split the sample into two subsamples according to the average annual number of employees in the company. Secondly we aim at rigidities in manufacturing and services.

We distinguish between small and large companies. Small companies are those, which have up to 40 employees. On the other hand, large companies have at least 90 employees. Thresholds 40 and 90 employees were set in order to split the sample into three subsamples with similar number of observations. Table 3.8 reports the results. Since we did not find statistically significant presence of rigidity we can conclude that small employers can better adjust wage costs according to their needs. On the other hand, we found significant nominal wage rigidities in larger companies in most of the years of the period studied (from 2004 up to 2007).

Table 3.8: Nominal wage rigidity according to size of the company

	Small (<40 empl.)			Large (>90 empl.)		
Years	fwcp sim	nwr	p-val	fwcp sim	nwr	p-val
2007/2006	0.037	0.009	0.181	0.152	0.025	0.003**
2006/2005	0.038	0.010	0.167	0.150	0.029	0.001**
2005/2004	0.004	0.001	0.467	0.129	0.030	0.003**
2004/2003	0.007	0.002	0.434	0.077	0.021	0.031*
2003/2002	-0.055	NA	0.944	0.014	0.003	0.391
2002/2001						
2001/2000	-0.021	NA	0.730	0.115	0.022	0.011*

Source: authors' calculation.
** (*) Note: DRWR are statistically significant at 1% (5 %) level of significance.
Note: fwcp_sim – simulated fraction of wage cuts prevented, nwr – share of companies affected by nominal rigidity, p-val – p value.

Further, we divided the sample according to economic activity. Here we report the results only for manufacturing and service (Table 3.9). It turns out that companies in the service sector can better adjust wage costs according to their needs, whereas manufacturing seems to be more rigid in wage formation.

Table 3.9: Nominal wage rigidity according in manufacturing and services

	Manufacturing			Services		
Years	fwcp sim	nwr	p-val	fwcp sim	nwr	p-val
2007/2006	0.092	0.018	0.023*	0.072	0.020	0.027*
2006/2005	0.171	0.035	0.000**	0.009	0.002	0.419
2005/2004	0.073	0.017	0.041*	0.021	0.007	0.284
2004/2003	0.042	0.012	0.135	-0.002	NA	0.527
2003/2002	0.015	0.003	0.374	-0.040	NA	0.872
2002/2001						
2001/2000	-0.013	NA	0.630	0.023	0.007	0.289

Source: authors' calculation.
** (*) Note: DRWR are statistically significant at 1% (5 %) level of significance.
Note: fwcp_sim – simulated fraction of wage cuts prevented, nwr – share of companies affected by nominal rigidity, p-val – p value.

3.4 Conclusions

Having reproduced a histogram location approach on the industrial level, we may conclude as follows:

From the accessible data of recent years, it is relevant to use histogram location approach and thus search for downward nominal wage rigidities in EU-15 and in EU-10 countries separately. Unification of all observations into one full sample (of EU-15 and EU-10 countries) may be a subject to distortion originating in specific features of data in the two subgroups. Extent of distortion however may be quantified and therefore interpretation of unified sample is possible with caution.

Nominal wages are rigid downward especially in Austria, France, Belgium and the Netherlands with wage cuts prevented in excess of 40%. For all the other countries of the EU-15 sample we have not found significant wage rigidities, even though we found over 20% downward wage rigidities in U.K. and Finland. These findings are generally in line with other results from cross-country studies. Further it suggests that decreasing trend of downward nominal wage rigidities in time identified in Holden and Wulfsberg (2007) experienced its bottom point in the 1990s´ while since trending upwards again.

Nominal wages in new EU member states are relatively flexible all across the countries we have included in the sample. In case of Slovakia and Slovenia however, final result cannot be drawn. Having too few negative observations in the sample, there is higher sensitivity to random disturbances, which makes such results difficult to interpret.

In the second part of the paper we employ histogram location approach on company level data in Slovakia. The modification of this paper is the adoption of the methodology proposed by Holden and Wulfsberg (2007) to a company level data. The data sample used covers hourly compensations in the time period between 2000 and 2007. The estimated extent of both nominal and real rigidity is relatively small. Conclusion that total compensations are rather flexible supports the decision of euro adoption in 2009.

We identify nominal wage rigidity only in the second part of the observed period (2005-2007). Although the methodology allows us to estimate lower bound of wage rigidity, based on estimated figures we can conclude that downward wage rigidity is small in the Slovak Republic. The computed share of companies affected by nominal wage rigidity ranges from 1.5 % in 2005 to 2.2 % in 2006 As a result, companies paid almost 300 million Sk (estimated number) more due to nominal wage rigidities in 2007. In macroeconomic sense this makes additional 0.14 percentage point of wage growth, which is a negligible effect. According to the methodology used, the extent of real wage rigidity is comparable to the degree of nominal wage rigidity and ranges between 2.1 % and 2.4 %. Detailed analysis shows that small companies can better adjust wage costs according to their needs. On the other hand, we found significant nominal wage rigidities in larger companies in most of the years in the period studied. We can also conclude that companies in the service sector can better adjust wage costs according to their needs whereas manufacturing seems to be more rigid in wage formation.

Appendix

Construction of the notional distribution and measurement of the downward wage rigidity

In this part we briefly introduce the methodology which helps us to identify measure and test the extent of prevented wage cuts. Detailed description can be found in Holden and Wulfsberg (2007).

The main assumption of this approach is that absence of rigidity is present in some country years[47] in the sample. Thus the first task is to choose those country years which represents the environment where wage rigidity doesn't bind. We decided to pick those with the highest nominal and real median wage growth within the sample. Selected wage change distributions are normalized by subtracting the corresponding medians and dividing by standard deviations (3.1) . The resulting wage change distribution is called the hypothetical distribution and can be described in the following mathematical notation:

$$\Delta w^{hyp} = \left(\frac{\Delta w_{jit} - med_{it}}{\sigma_{it}} \right), \tag{3.1}$$

where index *j* stands for industry or firm, *i* is a symbol of the country and *t* denotes year.

The hypothetical distribution is used to construct the notional distribution for each country year which represents the hypothesis of no rigidity. Therefore we multiply the common hypothetical distribution by corresponding standard error and then we add the country year median (3.2). The notional wage changes distribution is the constructed as follows:

$$\Delta w_{it}^{not} = \Delta w^{hyp} * \sigma_{it} + med_{it} . \tag{3.2}$$

The notional incidence rate is a share of the number of industries/firms that are supposed to cut wages (according to notional assumption of no rigidity) to the total number of firms.

$$q_{it}\_not = \frac{\#\left(\Delta w_{it}^{not} < 0\right)}{\#\Delta w_{it}^{not}} . \tag{3.3}$$

The empirical incidence rate is computed similarly:

$$q_{it}\_emp = \frac{\#\left(\Delta w_{it}^{emp} < 0\right)}{\#\Delta w_{it}^{emp}} . \tag{3.4}$$

[47] Note: We use the term country year. However, in the part 4.2 it stands for one year as we focus only on Slovakia.

The extent of rigidity is measured by a comparison of the amount of negative empirical and notional wage changes, represented by the incidence rates. Thus we are interested in the FWCP index:

$$FWCP\_not = 1 - \frac{q_{it}\_emp}{q_{it}\_not}. \tag{3.5}$$

FWCP reflects the share of industries/firms which didn't realize the wage cut thought they were supposed to cut wages to the total number of industries/firms supposed to cut wages in the no rigidity environment.

Testing of significance

In order to test the significance of computed shares of the industries/firms affected by wage rigidity we conduct the following test/procedure. The null hypothesis is that the extent of wage rigidity is statistically insignificant (no rigidity in wages).

We employ the test to examine the significance of rigidity procedure in part 3.3 for every country and for the whole sample and in part 3.3.2 for every year.

We use the binomial test however in a slightly simplified version. Instead of computing the exact probabilities we rather simulate draws from binomial distribution 5 000 times. Such a simplification dramatically decreased the computational requirements. It is important to add that part of the results were double-checked and normal approximation was used for the Binomial distribution. Both tests gave us the same results.

First step is to draw from the standard binomial distribution *B(n,p)* n times, where *n* stands for the number of trials and *p* is a success probability. Particularly, in this context, *n* is a number of empirical observations belonging to the respective country year and *p* is the notional incidence rate (q_{it}_not). We proceed by repeating this step 5 000 times. Afterwards we compute the average number of successive draws: $S^{it} = \frac{1}{5000}\sum_{k=1}^{5000} S_k^{it}$.

Dividing S^{it} by *n* we obtain simulated incidence rate (q_{it}_sim) and also are able to compute belonging fraction of wage cuts prevented ($fwcp_{it}\_sim$). [48] We then count the simulated amount of wage cuts S^{it} higher than the number of observed cuts in the corresponding empirical distribution. We call it H^{it}. Finally, the p-value is obtained as a $p = 1 - \frac{H^{it}}{5000}$. Based on the p-values, if the p-value is smaller than the significance level, one rejects the null hypothesis of no rigidity.

[48] Note: $fwcp_{it}\_sim$ is very similar to $fwcp_{it}\_not$ by construction. The higher is the number of simulations, the more these two values converge.

Summary

Using household expenditure surveys in Slovakia, we demonstrate significant differences between cohorts entering the labor market before and after 1990. On the one hand, returns to human capital are lower for education acquired before market reforms. On the other hand, the early labor market cohorts enjoyed also easier access to housing. We find that both effects seem to counteract each other to a significant degree. Older employees face lower returns to human capital, which lowers their disposable income. Keeping other effects unchanged, this would result in negative implications on their welfare. However, older households enjoyed also a preferential access to housing. We find that both effects seem to counteract each other to a significant degree. Older employees face lower returns to human capital, which lowers their disposable income. Keeping other effects unchanged, this would result in negative implications on their welfare. However, older households enjoyed also a preferential access to housing. Although the effects are difficult to quantify exactly, the magnitude of both effects, and their expected variability among individuals, lead to conclusion that it is difficult to identify winners and losers of transition, at least in the example of Slovakia.

In the second chapter, we look at the impact of disposable income and financial wealth on the household consumption from the macro perspective. We show that financial wealth influences the consumption and has lower elasticity than disposable income. We also look for appropriate proxy for financial wealth from four candidates. It turns out that the best proxy is a sum of monetary aggregate M2 that represents a significant part of household portfolio and assets invested in mutual funds. Moreover, we examine the impact of other relevant variables on consumption. We conclude that the real consumption does not significantly respond to the change in interest rates. Based on a comparison of in sample and out of sample forecasts we propose a vector error correction model for consumption forecasting.

The third chapter presents an analysis of wage rigidity. It turns out that the nominal wages are rigid downward especially in Austria, France, Belgium and the Netherlands with wage

cuts prevented in excess of 40%. For all the other countries of the EU-15 sample we have not found significant wage rigidities. These findings are generally in line with other results from cross-country studies. Nominal wages in new EU member states are relatively flexible all across the countries in our sample. In case of Slovakia and Slovenia however, final result cannot be drawn. Therefore, we turn our attention to company level data in Slovakia. The estimated extent of both nominal and real rigidity is relatively small. Conclusion that total compensations are rather flexible supports the decision of euro adoption in 2009. We find nominal wage rigidity only in the second part of the observed period (2005-2007). Although the methodology allows us to estimate lower bound of wage rigidity, we can conclude that downward wage rigidity is small in the Slovak Republic. The share of companies affected by nominal wage rigidity ranges from 1.5 percent in 2005 to 2.2 percent in 2006. Not surprisingly, the detailed analysis shows that small companies can better adjust wage costs according to their needs. On the other hand, we find significant nominal wage rigidities in larger companies in most of the years in the period studied. We can also conclude that companies in the service sector can better adjust wage costs according to their needs whereas manufacturing seems to be more rigid in wage formation.

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